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Trust in the Central Bank and Inflation Expectations*

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Using micro data from the 2015 Dutch CentERpanel, we examine whether trust in the European Central Bank (ECB) influences individuals' expectations and uncertainty about future inflation, and whether it anchors inflation expectations. We find that higher trust in the ECB lowers inflation expectations on average and significantly reduces uncertainty about future inflation. Moreover, results from quantile regressions suggest that trusting the ECB increases (lowers) inflation expectations when the latter are below (above) the ECB's inflation target. These findings hold after controlling for people's knowledge about the objectives of the ECB.

JEL Codes: D12, D81, E03, E40, E58.

1. Introduction

For central banks, the management of economic expectations has become a key tool in conducting monetary policy (Blinder et al. 2008). To effectively manage expectations, a high level of public trust

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in central banks is necessary. For instance, a high level of trust in the commitment and ability of the European Central Bank (ECB) to keep inflation below, but close to, 2 percent can anchor medium- and long-term inflation expectations on this target and make the public view deviations from the target as temporary ones. More generally, a trustworthy central bank is more likely to increase public confidence about future price stability and the prospects of the economy, thereby boosting economic growth.[1]

A number of recent studies document that inflation expectations feed into important household financial decisions. Malmendier and Nagel (2016) show that households expecting higher inflation are less likely to invest in long-term bonds and more likely to borrow through fixed-rate mortgages compared with households expecting low inflation. Armantier et al. (2015) conduct a financially incentivized investment experiment and find that individuals' reported inflation expectations influence their investment choices in a way that is consistent with economic theory. Moreover, D'Acunto, Hoang, and Weber (2016) show that an increase in inflation expectations implies a higher readiness to purchase durable goods.

To the best of our knowledge, existing literature does not provide empirical evidence on the relationship between trust in central banks and inflation expectations.[2] Our paper fills this gap by investigating whether citizens' trust in the ECB contributes to individuals' expectations and uncertainty about price growth and the extent to which these expectations are anchored at the ECB's medium-term inflation target of below, but close to, 2 percent. More broadly, our

[1]For instance, persistently high inflation has been historically linked to increased uncertainty about price evolution and low public confidence in the economy's prospects. Such an environment has typically adverse effects on individuals' saving, consumption and investment decisions. As Bernanke (2013) states: "Expectations matter so much that a central bank may be able to help make policy more effective by working to shape those expectations."

[2]The relation between trust and inflation expectations has been recently recognized as a highly policy relevant topic that merits empirical investigation: "Another field in which public trust in central banks might prove important is for the understanding of the formation of household inflation expectations [. . . .] If low public trust in central banks is associated with higher household inflation expectations, then swings in public trust in ECB also directly affect its ability to deliver on its mandate, although the empirical relevance of this proposition has yet to be tested" (Ehrmann, Soudan, and Stracca 2013, pp. 782–83).

paper contributes to existing literature which finds that higher social capital and trust are conducive to better economic outcomes such as higher growth (see, e.g., Zak and Knack 2001; Tabellini 2010).

We use recent micro data from the CentERpanel, a representative survey of the Dutch-speaking population in the Netherlands, sponsored by the Dutch National Bank (DNB). We survey individuals during the first half of 2015 and ask them a set of specially designed questions that allow us to construct individual-specific measures of expected inflation and inflation uncertainty. Given that the ECB's policy impacts a broad range of economic outcomes, we collect similar information on expectations regarding economic growth. We ask also how much individuals trust the ECB. Similar trust questions that aim to measure public trust in the ECB and in other European institutions are regularly asked in Eurobarometer surveys since the early 2000s. Existing literature has used Eurobarometer data to examine possible determinants of the ECB trust (Ehrmann, Soudan, and Stracca 2013; Bursian and Fürth 2015). Instead, our paper explores the influence of trust in the ECB on inflation expectations. Finally, the survey contains a series of questions about the objectives of the ECB that allow us to distinguish whether reported trust in the ECB reflects the perceived credibility of the institution or simply knowledge about its role.

Our analysis offers a number of novel findings. First, higher trust in the ECB induces, on average, lower one-year-ahead inflation expectations. This relationship, however, is not uniform across different percentiles of the distribution of inflation expectations. Second, higher trust contributes significantly to lower individual uncertainty about future price growth, thus implying a form of anchoring of inflation expectations. Third, higher trust in the ECB is associated with higher inflation expectations when the latter are at the lower end of the sample distribution, while the opposite is true when people have inflation expectations at the upper end. This effect is particularly strong for those who report inflation expectations above the ECB target, while it is less robust for the part of the distribution that corresponds to expected inflation that is below, but close to, 2 percent.

Taken together, these results point to the role that trust in the ECB can play in anchoring consumers' inflation expectations around the ECB's medium-term inflation target. If expectations are well

anchored around the target, the public should be confident about its inflation estimate and react little to short movements of higher- or lower-than-expected inflation. Moreover, if anchoring of public expectations occurs close to the central bank's inflation target, equilibrium prices should adjust faster toward this target (Bernanke 2013).

As the survey was conducted in the first half of 2015, it is noteworthy that the estimated effect of trust on anchoring inflation expectations is economically important even in an environment of low interest rates and inflation. This suggests that factors such as citizens' trust in the central bank can be important when conventional monetary policy tools turn out to be least effective.

Fourth, we examine whether estimated effects of trust in the ECB reflect public knowledge about the ECB objectives per se (e.g., the numerical target for inflation) or credibility more broadly defined (e.g., in ECB's commitment to maintain price stability). To this end, we control for respondents' knowledge regarding the ECB's objectives or their financial sophistication, and results remain unchanged. This suggests that people's perceptions about the credibility of an institution can influence their inflation expectations over and above their knowledge about the specific objectives of the institution. It also highlights the long-term advantages of establishing a reputation for central bank credibility, as it could operate beyond precise knowledge about central bank's objectives or temporary deviations from them.[3] Fifth, trust in the ECB is also positively associated with expectations about economic growth, but not with the expected variability of output growth.

In the investigation, we estimate several empirical models to make sure that the effects we uncover are due to individuals' trust in the ECB as an institution and not to other possible confounding factors. As we discuss in detail in section 4, we identify our parameters of interest through instrumental-variable (IV) estimation using information on episodes of cheating by repair persons that respondents have experienced in the past few years. We assume that exposure to such events is correlated with the social capital component of trust in the ECB, but has no independent effect on inflation expectations.

[3]Kril, Leiser, and Spivak (2016) provide a thorough discussion of the economic advantages of central bank credibility.

We also use, as a second instrument, the trust that respondents have in other people, as interpersonal trust is unlikely to directly shape inflation expectations. The test of the overidentification restrictions strongly suggests that the null hypothesis of the exogeneity of these instruments cannot be rejected.

Moreover, we control not only for standard socioeconomic characteristics but also for respondents' general economic knowledge by means of three standard questions that have been developed by Lusardi and Mitchell (2014) and are widely used to measure financial literacy. In a related vein, we measure and control for individuals' specialized knowledge about the ECB's objectives. The main findings on the role of trust in the ECB remain unaffected when we control for these indicators of knowledge about the ECB and financial literacy, which suggests that genuine trust in the institution affects inflation expectations over and above knowledge about the ECB's mission or economics in general. Furthermore, our main findings are unaffected when a measure of individual optimism is included in our specifications.

The remainder of the paper is organized as follows. Section 2 reviews relevant studies on inflation expectations and trust in institutions. Section 3 describes the survey data. Section 4 presents the baseline empirical results on inflation expectations, uncertainty, and anchoring. Section 5 presents a number of robustness checks, and section 6 contains additional results on expectations about economic growth and output variability. Section 7 summarizes our main findings.

2. Inflation Expectations and Trust

Our paper spans different strands of the literature on inflation expectations and trust. Recent studies explore links between survey-based inflation expectations, anchoring around central banks' inflation target and understanding of monetary policy operations. In particular, Kumar et al. (2015) survey firms' managers in New Zealand and find evidence that their inflation expectations are not anchored, despite the Reserve Bank's inflation targeting for more than 25 years. Carvalho and Nechio (2014) provide evidence that some U.S. households form expectations consistent with a Taylor rule. Kril, Leiser, and Spivak (2016) use survey data from Israel to examine the determinants

of central bank credibility and trust. Based on a detailed set of questions, they document that the public perception about central bank credibility is primarily linked to views regarding the professionalism and independence of the central bank and not with its transparency per se.

While there are only a few studies using survey-based information on inflation expectations, there are several papers on anchoring and inflation expectations based upon financial market instruments such as inflation options, swaps, and index-linked securities. Some of these studies investigate central bank credibility, mostly after estimating it using financial market-based expectations (see, e.g., Gürkaynak, Levin, and Swanson 2010; Gerlach-Kristen and Moessner 2014). Kril, Leiser, and Spivak (2016), on the other hand, argue that these inflation expectations reflect the combined assessment of economic conditions and central bank credibility. Instead, they measure central bank credibility based upon the confidence of the public in central bank forecasts for inflation and economic growth. Their discussion and results indicate that while trust and credibility are multidimensional concepts, they are closely related to each other.[4] In our paper, we interpret trust as a measure of the institutional credibility of the ECB, while in a robustness analysis we show that such institutional trust is not confounded by knowledge of the ECB's objectives.[5]

A related group of studies focuses on the role of central bank communication for financial market outcomes; Blinder et al. (2008) provide a thorough review. For example, Ehrmann and Fratzscher (2005) show that press statements by central banks have an immediate impact on financial markets and also affect the latter's ability to anticipate future monetary policy decisions. In contrast to these studies, we elicit survey-based consumer expectations on future

[4]Indeed, the concept of central bank credibility was popularized as a solution to the time-inconsistency problem discussed by Kydland and Prescott (1977). One solution to this problem was to delegate monetary policy to a central bank or another institute with a high perceived credibility so that the public can have sufficient trust in the central banks' ability to withstand the temptation to create surprise short-term inflation.

[5]In the terminology of Kril, Leiser, and Spivak (2016), the knowledge component refers to reasoned trust/credibility while institutional trust refers to an affective form of trust/credibility.

inflation and its uncertainty and link them to reported trust in the ECB.

There is also a large body of literature studying the implications of trust in other people and social capital in different domains: economic growth (Zak and Knack 2001; Guiso, Sapienza, and Zingales 2004; Tabellini 2010), confidence in the government (Knack and Keefer 1997), financial integration (Ekinci, Kalemli-Özcan, and Sorensen 2009), cross-country trade (Guiso, Sapienza, and Zingales 2008), and household portfolios (Guiso, Sapienza, and Zingales 2008; Georgarakos and Pasini 2011). Another group of studies explores the determinants of trust in institutions, and central banks in particular.[6] A robust finding of the literature on social capital is that trust in other people changes slowly over time, given that social capital consists of a large inherited component of social values and norms (Tabellini 2010).[7] On the other hand, Stevenson and Wolfers (2011) point out that specific trust in financial institutions is more responsive to prevailing economic conditions.

3. The CentER Internet Panel

We use data from the CentER Internet panel, which is sponsored by DNB and maintained by CentERdata at Tilburg University.[8] The baseline survey is conducted annually, and collects detailed information on a range of demographic and economic variables for a representative sample of Dutch-speaking households. In addition to the baseline survey, respondents answer questions during the course of a year in special-purpose surveys.

We designed such a special-purpose survey that measures individuals' expectations and uncertainty about future price growth as

[6]For recent studies see Wälti (2012), Ehrmann, Soudan, and Stracca (2013), Bursian and Fürth (2015), or van der Cruijsen, de Haan, and Jansen (2016).

[7]Societal attitudes are often considered as slow moving, but there are also instances in which they adapt relatively fast to the environment. For example, Guiso, Sapienza, and Zingales (2006) refer to culture as those customary beliefs and values that are transmitted fairly unchanged from generation to generation. On the other hand, Fernandez (2011) mentions as a counterexample the rapid change (i.e., within half a century) in attitudes toward married women working.

[8]Panel members are recruited via personal or telephone interviews. If, after being selected for panel participation, it turns out that respondents have no computer with Internet access, CentERdata provides them the necessary equipment.

well as trust in the ECB. We administered the special survey to every panel participant aged 18 and older in January 2015. The survey was repeated in June 2015 to account for a possible seasonal pattern in responses and to increase sample size.

To elicit the distribution of expected inflation, we follow a similar procedure as in Guiso, Jappelli, and Pistaferri (2002), Guiso, Jappelli, and Padula (2013), and Christelis et al. (2020), who estimate the subjective distribution of expected income, pension replacement rate, and consumption, respectively. Specifically, we ask respondents to report the minimum (y_m) and the maximum (y_M) values of percentage change in the level of prices 12 months ahead. Subsequently, we ask them to indicate on a 0–100 scale the probability that the average change in prices in the next 12 months will be higher than the midpoint between the reported minimum and maximum, i.e., $\pi = Prob\left(y > \left(y_m + y_M\right)/2\right)$. The questions we use are reported in section A.1 of the online appendix available at http://www.ijcb.org.

To estimate the moments of the subjective distribution of expected inflation, we rely on the assumptions and methods used by Guiso, Jappelli, and Pistaferri (2002) and Christelis et al. (2020) to estimate the subjective distribution of future income and consumption, respectively. We assume that the subjective distribution is either simple triangular (i.e., symmetric around $\left(y_m + y_M\right)/2$, assuming $\pi = 0.5$) or split triangular ($\pi \neq 0.5$; see figure A.1 in the online appendix). Based on the elicited values of y_m and y_M (and of π if we assume a split triangular distribution), we compute the individual-specific mean and standard deviation of the distribution of expected inflation one year ahead. The formulas of these statistics are reported in section A.2 of the online appendix.[9]

We set values of the moments of the individual-specific subjective expected inflation distributions to missing when y_m, y_M, or π are missing, or when respondents choose the "don't know" option. The original sample includes 4,333 observations in the two survey waves. Due to missing values, the estimation sample includes 3,117

[9]We assume that y_m and y_M represent the actual minimum and maximum of the distribution. This is potentially a strong assumption. Dominitz and Manski (1997) use the percentage change format to elicit the subjective income distribution, and show that individuals associate the "lowest possible" (and "highest possible") with low (respectively, high) probability.

observations that allow the computation of inflation expectations using the simple triangular distribution, and 3,019 observations using the split triangular distribution.

The survey also asks individuals to indicate their level of trust in the ECB on a 0 to 10 scale, where 0 denotes no trust in the ECB whatsoever, while 10 denotes full trust. A similar question has been regularly asked in Eurobarometer surveys since the early 2000s so as to measure public trust in the ECB as well as in other institutions such as the United Nations, the European Union, the national Parliament, and the national government.[10] Eurobarometer data have been used by several studies to examine determinants of trust in the ECB or its evolution over time (see, e.g., Wälti 2012; Ehrmann, Soudan, and Stracca 2013; Bursian and Fürth 2015).

It is important to note that we ask the question on trust in the ECB without making any explicit reference to its objective of an inflation rate at or slightly below 2 percent. We do this for a number of reasons. First, we want to avoid conditioning the answers to the question on trust in the ECB. Given respondents' expected inflation, reminding them of the ECB's target inflation rate may change their answer. For instance, those who expect inflation next year to be 3 percent, when reminded that the ECB target rate is 2 percent, could reduce their reported trust. Second, we need to measure respondents' trust in the ECB given their current knowledge about economic affairs, without influencing this knowledge. Third, we aim to distinguish the notion of trust as institutional credibility from that of trust as institutional knowledge. This is why we control separately for the latter in our empirical specification, as discussed in section 5 below. Making this distinction would have been impossible if we had mentioned explicitly the ECB inflation target in the question. Finally, while the primary target of the ECB is the inflation rate, its mandate states that it also aims to promote economic growth, with no prejudice to achieving its inflation target. The weight given to the two objectives can vary over time in practice.

[10]The Eurobarometer question has three answer options: "tend to trust," "tend not to trust," or "do not know." Our question asks for the intensity of trust using a 0–10 scale (similar to the answer scale in questions on trust in other people used in surveys such as the European Social Survey).

To separate the notions of trust and institutional knowledge, we ask a series of questions to measure knowledge about the ECB's objectives and basic financial literacy (see section A.1 of the online appendix for the exact wording of these questions). Other studies have asked related questions that capture knowledge of the ECB's objectives. For instance, van der Cruijsen, Jansen, and de Haan (2015) find a low prevalence of knowledge of the ECB's objectives. In our survey, we present six statements about the ECB's objectives. These statements mention specific numerical targets (e.g., for unemployment), in order to make sure that individuals are not confused with the fact that the ECB's policies can have broader positive economic consequences beyond price stability. Importantly, to avoid framing effects, we place the questions on expected inflation, trust in the ECB, and knowledge about its objectives in separate sections of the questionnaire.[11]

To purge the relation between trust and inflation expectations from the effect of financial literacy, we ask the three basic financial literacy questions proposed by Lusardi and Mitchell (2014). The questions aim to capture individuals' numeracy and understanding of basic economic concepts such as interest rates, inflation, and risk diversification, and have been used in many studies and countries (see Lusardi and Mitchell 2014 for an overview).

Figures 1 and 2 show the distribution of the expected minimum and maximum levels of inflation 12 months ahead.[12] For each observation in the sample, the maximum is greater than the minimum. Figure 3 illustrates the distribution of the probability that the expected inflation is above the midpoint of the expected minimum and maximum values. As can be seen, there is a prevalence of "50 percent" responses but also a sizable number

[11]As the latter set of questions makes reference to numerical targets that may (or may not) be among the objectives of the ECB, we have placed them in the last part of the questionnaire to avoid bias in the answers to questions on inflation expectations and/or trust in the ECB.

[12]Note that while our survey focuses on 12-months-ahead inflation expectations, in recent consumer expectations surveys, respondents seem to provide similar answers to questions about short-term (one-year) and medium-term (three-year) inflation expectations (see, e.g., the New York Federal Reserve Bank's online reports of their survey on consumer expectations: https://www.newyorkfed.org/microeconomics/sce#indicators/inflation-expectations/g1).

Figure 1. Histogram of the Minimum Expected Inflation Level

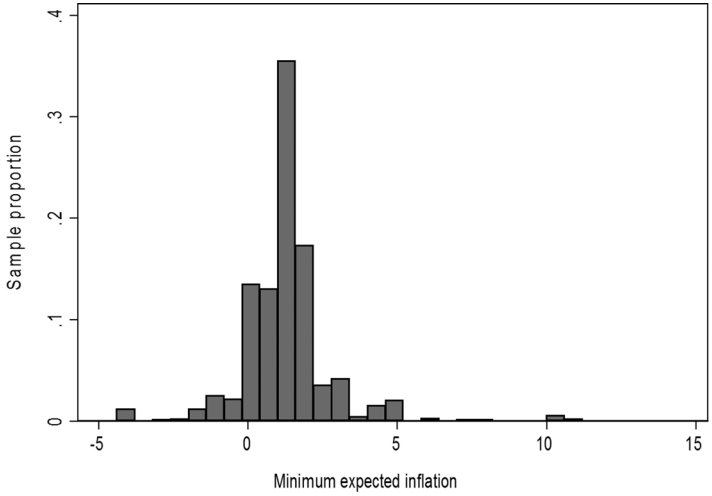


Figure 2. Histogram of the Maximum Expected Inflation Level

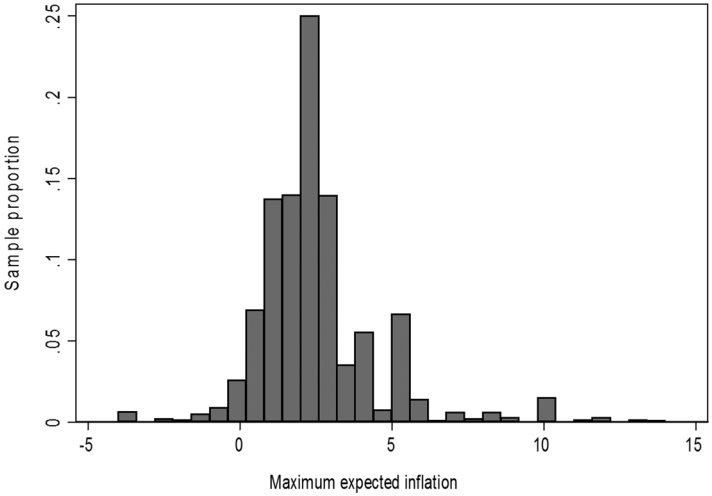
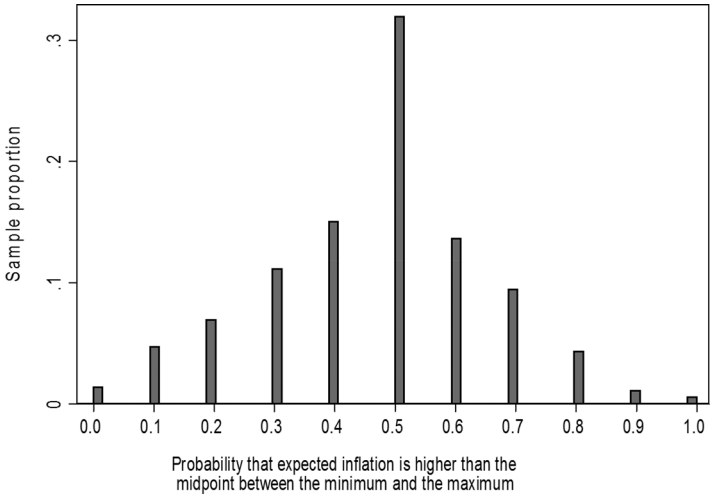


Figure 3. Histogram of the Probability that Expected Inflation Is Above the Average of the Expected Minimum and Maximum Values



of responses with values larger or smaller than 50 percent. We note that responses to the question on this probability, which is arguably more difficult to answer, are not used in our baseline regressions, which rely on expectations computed using the simple triangular distribution.

Table 1 reports cross-sectional estimates of the median and coefficient of variation of the subjective distributions of expected inflation and expected growth, assuming that the underlying distribution is simple (i.e., symmetric) triangular. At the median, the minimum expected inflation is 1 percent, while the maximum is 2 percent (the means are equal to 1.3 percent and 2.7 percent, respectively). The median probability is 0.50 (average 0.47). Assuming that the distribution is simple triangular, we estimate that the sample median of expected inflation is 1.6 percent (average 2 percent), while the median (mean) variance of the distribution of expected inflation is 0.04 percent (0.21 percent). The coefficient of variation of the distribution of expected inflation—that is, the ratio of its standard deviation to its mean—is 1.09, which implies that respondents experience considerable uncertainty about future inflation.

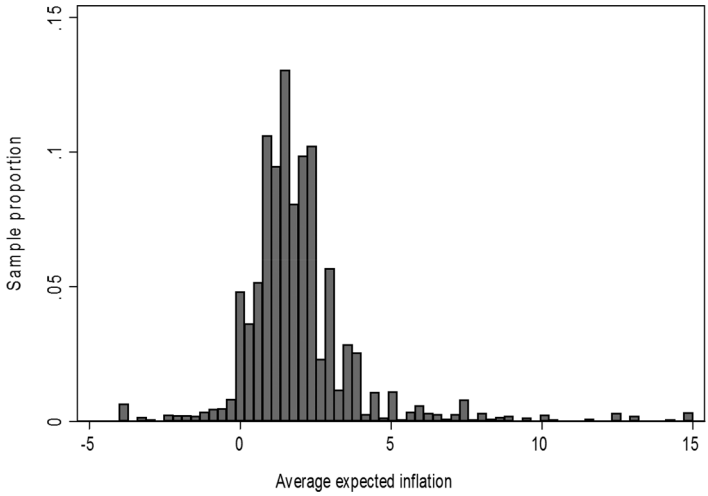
Table 1. Descriptive Statistics

Variable	Mean	Median	Std. Deviation
Minimum Expected Inflation Level	1.303	1.000	1.889
Maximum Expected Inflation Level	2.699	2.000	2.743
Probability that the Expected Inflation Level is Above the Average of the Expected Minimum and Maximum Values	0.469	0.500	0.185
Expected Inflation	2.001	1.600	2.187
Variance of Expected Inflation	0.208	0.042	0.838
Minimum Expected Growth Rate	0.965	1.000	1.499
Maximum Expected Growth Rate	1.967	2.000	1.782
Probability that the Expected Growth Rate is Above the Average of the Expected Minimum and Maximum Values	0.415	0.500	0.186
Expected Growth Rate	1.465	1.500	1.520
Variance of Expected Growth Rate	0.108	0.020	0.575
Trust in the ECB	4.769	5.0	2.163
Age	50.8	51.0	16.6
Female Householder	0.452	0.0	0.498
Household Size	2.488	2.0	1.252
Has a Spouse/Partner	0.747	1.0	0.435
High School Graduate	0.383	0.0	0.486
University Graduate	0.309	0.0	0.462
Household Income	2,721.8	2,450.0	2,004.8
Number of Observations	3,117	3,117	3,117

Note: All magnitudes related to the expected inflation and growth rate are expressed in percentage points.

Given that inflation in the Netherlands in 2015 and 2016 was extremely low by historical standards (0.2 and 0.1 percent according to the Harmonised Index of Consumer Prices), the median expected inflation (1.6 percent) turned out to be an overestimate of the realized values. This is in line with evidence from other surveys on inflation expectations (see Coibion and Gorodnichenko 2015 and Kliesen 2015 for evidence from the University of Michigan Survey of Consumers) and could also reflect perceptions of price developments in the particular bundle of goods that each household regularly spends on.

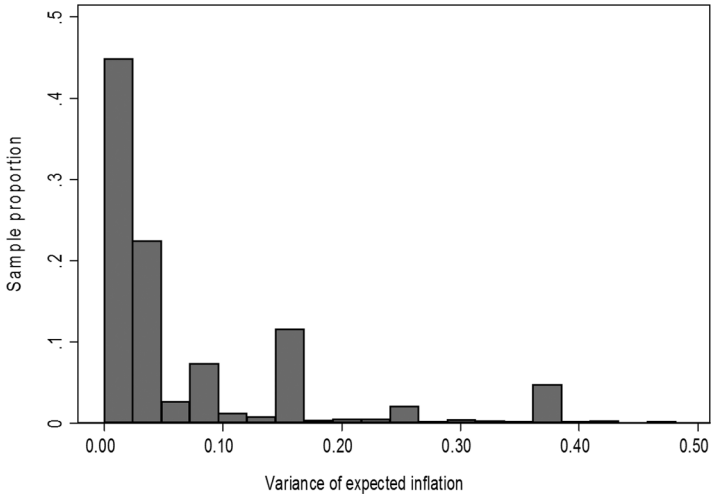
Figure 4. Histogram of the Mean of Expected Inflation



The respective sample median (mean) of expected gross domestic product (GDP) growth is 1.5 (1.47) percent, while the median (mean) variance of the distribution of expected growth is 0.02 percent (0.11 percent). As regards trust in the ECB (asked on a 0 to 10 scale), the median (mean) is 5 (4.77) and the standard deviation is 2.16. The table also shows summary statistics on socioeconomic characteristics that are taken into account in the estimation (age, household size, marital status, education, and income).

Cross-sectional averages summarize the expected inflation distribution of a typical individual but hide important heterogeneity across individuals. Assuming that the underlying distribution is simple triangular, figures 4 and 5 plot the histogram of the means and variances, respectively, of the 3,117 individual-specific distributions of expected inflation. Both figures highlight the considerable heterogeneity in expected inflation distributions. For instance, for 6.7 percent of individuals the mean expected inflation is 0 or negative, for another 20.4 percent it is between 0 and 1 percent, for another 38.7 percent it is between 1 and 2 percent, and for the remaining 34.2 percent the mean expected inflation is higher than 2 percent. The cross-sectional distribution of variances is also heterogeneous, with roughly 13 percent of respondents exhibiting zero variance in

Figure 5. Histogram of the Variance of Expected Inflation



their expected inflation distribution (i.e., they do not report any uncertainty about future inflation).

4. Empirical Results

4.1 Regression Analysis

Before moving to regression analysis, we plot the mean of expected inflation and its variance by bins of trust in the ECB. The results are shown in figures 6 and 7, respectively. The figures suggest a negative association between trust in the ECB and both expected inflation and the variance of expected inflation.

Since the patterns of figures 6 and 7 may be influenced by other confounding variables, we estimate the relation between average expected inflation or the variance of expected inflation and reported trust in the ECB by estimating the following equation:

$$g_i\left(\pi_i^e\right) = \alpha + \beta trust\_ECB_i + \gamma X_i + \varepsilon_i, \tag{1}$$

where g is a function denoting either mean or variance of expected inflation π^e deduced from a simple triangular or split triangular distribution and taking individual-specific values. Variables in $\boldsymbol{X}$

Figure 6. Average Expected Inflation by Levels of Trust in the ECB

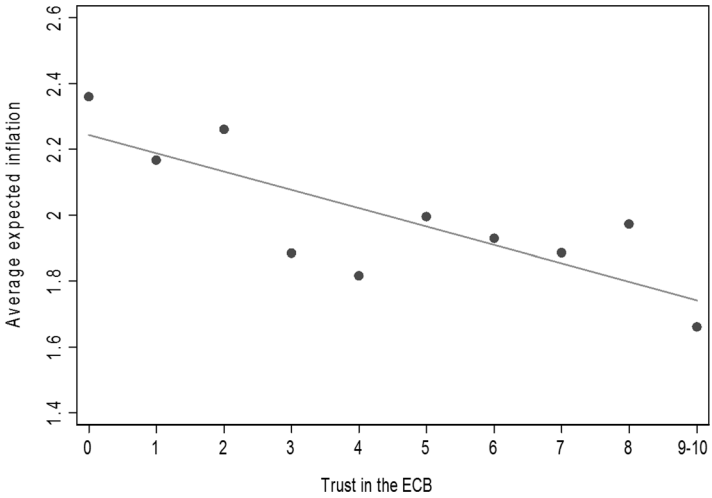
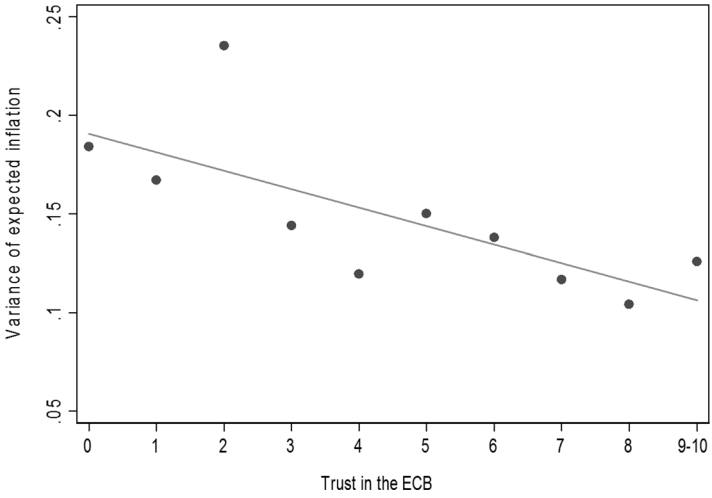


Figure 7. Variance of Expected Inflation by Levels of Trust in the ECB



include demographics, such as age (by means of a second-order polynomial) and gender of the respondent, whether (s)he has a partner, size of the household, whether the respondent is a high school graduate or has a college degree, and household income. Finally, we include a survey wave dummy to take into account country-wide conditions (e.g., actual past or current inflation and GDP growth) as well as region fixed effects that control for economic developments at the regional level.

To reduce the influence of outliers, we winsorize the mean and variance of expected inflation at the top and bottom 1 percent of the observations; that is, we set the values of those observations equal to those at the 99th and 1st percentiles, respectively. We also use Huber-White robust standard errors clustered at the household level to take into account that, for some households, multiple members can participate in the survey.

First, we estimate (1) for mean expected inflation using ordinary least squares (OLS). Results are shown in columns 1 and 2 of table 2. We find a negative association between trust in the ECB and average inflation expectations that is statistically significant at the 1 percent level. The OLS coefficient of trust in the ECB is equal to 0.055 percentage point, which implies that a one-standard-deviation increase in trust (equal to 2.16) is associated with a reduction in expected inflation of 0.12 percent, which corresponds to 7 percent of the sample average of expected inflation. As we show below, however, the average response obtained by OLS is not uniform across different percentiles of the expected inflation distribution.

Household size is negatively associated with higher inflation expectations.[13] This may reflect the fact that households with more than one member include several potential shoppers who may in turn be more efficient in identifying cheaper products and services, make better deals, or reduce per-unit cost by purchasing larger quantities. The coefficients of other demographic variables are not statistically different from zero.

Next, we use an IV approach to address endogeneity problems that may affect the OLS estimates of equation (1), and which could arise due to omitted variables and/or reverse causality. Finding valid

[13]Fritzer and Rumler (2015) report a similar finding for Austrian households.

Table 2. OLS and IV Results for Expected Inflation and Its Variance

	Average of Expected Inflation				Variance of Expected Inflation			
	OLS		IV		OLS		IV	
Variable	**Coeff. (1)**	**Std. Error (2)**	**Coeff. (3)**	**Std. Error (4)**	**Coeff. (5)**	**Std. Error (6)**	**Coeff. (7)**	**Std. Error (8)**
Trust in the ECB	−0.0552	0.0190***	−0.1736	0.0597***	−0.0120	0.0030***	−0.0210	0.0100**
Age	0.0155	0.0152	0.0090	0.0170	−0.0040	0.0030	−0.0040	0.0030
Age Squared	−0.0001	0.0001	−0.0001	0.0002	0.0000	0.0000	0.0000	0.0000
Female	0.0041	0.0717	−0.0591	0.0812	0.0170	0.0130	0.0090	0.0150
Couple	0.0922	0.1013	0.0726	0.1108	0.0080	0.0200	0.0090	0.0210
Household Size	−0.1331	0.0353***	−0.1505	0.0401***	−0.0170	0.0080**	−0.0220	0.0090**
High School Graduate	0.1093	0.0989	0.1851	0.1175	0.0010	0.0180	0.0130	0.0210
College Graduate	0.0976	0.0929	0.2550	0.1259**	0.0180	0.0170	0.0400	0.0210*
Logarithm of Household Net Income	0.0066	0.0343	0.0275	0.0340	−0.0080	0.0090	−0.0040	0.0100
Constant	1.5528	0.5116***	2.1168	0.6219***	0.4700	0.1210***	0.4880	0.1430***
Region/Wave Dummies	Yes		Yes		Yes		Yes	
First-Stage F-Test			86.433				86.433	
Endogeneity Test			3.971				0.974	
P-value			0.046				0.324	
Test of Overidentifying Restrictions			0.766				0.314	
P-value			0.381				0.575	
Observations	3,055		2,632		3,055		2,632	

Notes: This table shows OLS and IV estimation results from models using expected inflation and its variance (both measured in percentage points) as dependent variables. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively.

instruments in our context requires finding variables that correlate with trust in the ECB but do not have a direct association with inflation expectations. The latter condition precludes the use as instruments of variables that may be correlated with institutional or economic knowledge, such as own experience with retail banking services or credit card fraud, as these may have an independent impact on inflation expectation formation. Instead, we exploit variation in the component of trust in the ECB that relates to social capital. The latter typically reflects prevailing social values and tends to be resilient to temporary variations in financial conditions.

Respondents in our survey report the frequency they have been cheated by a plumber, builder, car mechanic, or other repair person over the past five years. Roughly one out of five respondents report having been cheated by a repair person at least once. The identifying assumption is that those who have been cheated tend to trust less, and that part of this mistrust carries over to the trust they show in institutions. Cheating experiences on some common everyday exchanges are arguably exogenous to own actions, so that instrument validity rests upon the assumption that exposure to such incidents does not have a direct impact (i.e., other than through trust in the ECB) on individuals' inflation expectations.

To increase the efficiency of our estimates and generate overidentifying restrictions, we also use as an instrument the reported trust in other people. Trust in other people has a strong intergenerational component that consists of inherited social norms. Thus, it is likely to respond less to contemporaneous economic and personal conditions than trust attached to financial institutions (see Tabellini 2010 and Stevenson and Wolfers 2011). Accordingly, we assume that general trust in other people is likely to influence inflation expectations only through institutional trust in the ECB.

The IV results are shown in columns 3 and 4 of table 2. The F-test statistic from the first-stage regressions is equal to about 86, which suggests that the instruments are quite strong (results are shown in online appendix table A.1). In addition, both instruments are correlated with trust in the ECB in an expected way. Moreover, a test of overidentifying restrictions (Hansen's J-test) fails to reject the null hypothesis of joint instrument validity with high confidence. Hence, we find no evidence of consistency problems in the IV estimates. Finally, a Hausman test of the endogeneity of the variable of

interest, i.e., trust in the ECB, has a p-value equal to 0.046, which implies that one can marginally reject the null of exogeneity of trust in the ECB at the 5 percent significance level, and thus that IV estimation is likely advisable.

The IV estimate of the effect of trust in the ECB on expected inflation is −0.17 percentage point, that is, somewhat larger in absolute value than the OLS one. The IV estimate implies that an increase in trust in the ECB of one standard deviation reduces expected inflation by 0.38 percentage point, which is 22 percent of the sample mean. The larger absolute value of the IV estimate compared with the OLS one could be due to the presence of unobserved variables that affect positively both trust in the ECB and inflation expectations, thus leading to an algebraically larger (i.e., less negative, but smaller in absolute value) OLS estimate. Such unobservable variables could include, for instance, expectations about unemployment or the general state of the economy. Alternatively, the difference between OLS and IV estimates could be due to the fact that in the presence of a heterogeneous effect of trust in the ECB this estimate represents a local average treatment effect, that is, the effect of trust in the ECB on inflation for those who change their trust in the ECB due to changes in the instrument values. On the other hand, the OLS estimate represents the overall average treatment effect.

Having examined the role of trust in the ECB on average inflation, we focus next on the role of trust on inflation uncertainty. Recall that the survey design allows us to deduce a measure of uncertainty about inflation that is individual specific. That is, we can estimate a version of equation (1) in which the dependent variable represents the individual-specific expected variance of inflation. As was the case for the expected inflation, its variance is calculated using a simple triangular distribution.

OLS estimates are reported in columns 5 and 6 in table 2. We find that a higher trust in the ECB leads to a lower variance of the expected inflation distribution. The effect is precisely estimated (p-value < 0.01) and implies that a one-standard-deviation increase in trust in the ECB reduces inflation uncertainty by about 30 percent of its sample mean.

The corresponding IV estimate (shown in columns 7–8 of table 2) is again precisely estimated (p-value < 0.05) and implies that a

one-standard-deviation increase in the ECB reduces inflation uncertainty by about 50 percent of its sample mean. Once more, the J-test of overidentifying restrictions clearly indicates that the null hypothesis of joint instrument validity cannot be refuted. The Hausman test for endogeneity suggests that the null cannot be rejected. Based on these results, the preferred estimates are those derived using standard OLS.

4.2 Quantile Regressions

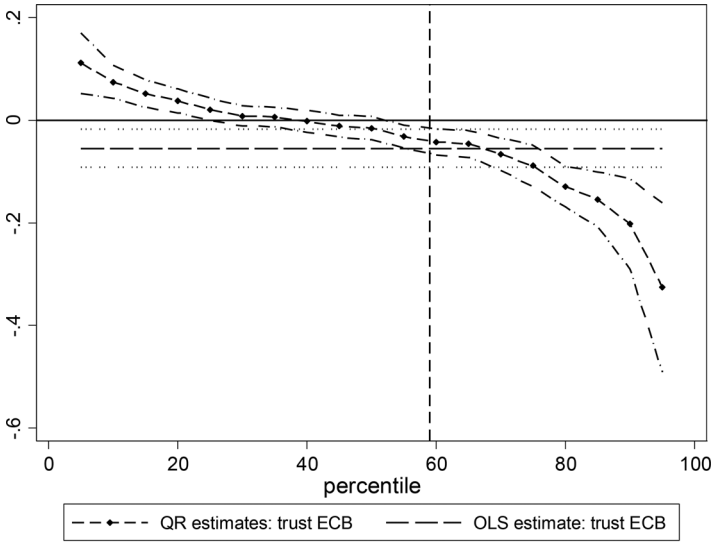
Results thus far suggest that higher trust in the ECB lowers inflation expectations on average. Nevertheless, this effect may not be symmetric across the distribution of expected inflation, that is, it might differ between those who have high and those who have low inflation expectations. In fact, the negative average estimated effect may simply reflect the public view that central banks are primarily concerned about actual inflation exceeding target inflation, and therefore are committed to raising interest rates to restrain inflation. That is, central banks have traditionally built their reputation as safeguards of price stability in situations when inflation tends to exceed their medium-term target. Consequently, a high level of public trust in the ECB could reflect trust in the ECB's commitment and ability to fight high inflation and thereby induces lower inflation expectations.[14]

Results also show that trust in the ECB significantly lowers inflation uncertainty, suggesting that trust may induce a form of anchoring of inflation expectations. We shed more light on this important policy issue by examining whether trust in the ECB contributes to anchoring of expectations around the ECB's target for an inflation below, but close to, 2 percent. To that effect, we estimate a series of quantile regressions (QR) to evaluate the effect of trust in the ECB across various percentiles of the expected inflation distribution.

Figure 8 plots the estimates and associated 95 percent confidence intervals of the effect of trust in the ECB derived from QR evaluated at every five quantiles of the conditional expected inflation

[14]The explanation that public trust in the ECB responds slowly to contemporaneous economic conditions is consistent with the important role of social capital in shaping trust in institutions (since social capital consists of a large, slow-moving, inherited component of social values and norms; see Tabellini 2010).

Figure 8. Quantile Regression Estimates for Expected Inflation



distribution (the respective estimates are reported in online appendix table A.2). The figure also plots the OLS estimate and confidence band from the baseline specification discussed above. The vertical line is drawn at the percentile of the expected inflation distribution that corresponds to a 2 percent inflation, that is, the ceiling of the medium-term ECB target.

The QR estimates imply that trust in the ECB raises inflation expectations among individuals with low inflation expectations. In particular, a one-standard-deviation increase in trust has a positive and strongly statistically significant effect at the 10th and 20th quantiles of the expected inflation distribution (corresponding to 0.35 percent and 0.75 percent expected inflation, respectively), which is equal to 0.16 percent and 0.08 percent, respectively. On the other hand, QR estimates at the sixth and higher deciles display a negative sign, suggesting a progressively stronger negative association of trust in the ECB with inflation expectations among those with higher inflation expectations. Notably, the magnitude of the QR coefficients is stronger on the upper part of the distribution compared with the lower one: a one-standard-deviation increase in trust

in the ECB decreases inflation expectations by 0.28 percent and 0.44 percent at the 80th and 90th percentiles of the expected inflation distribution, respectively.

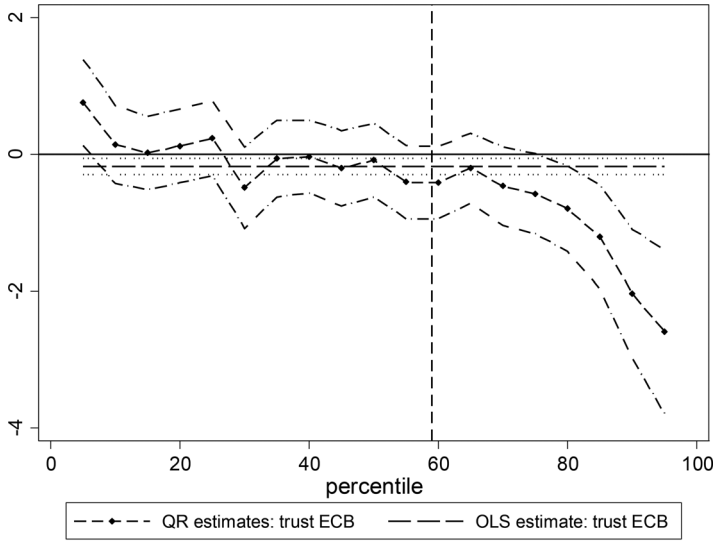
Finally, QR estimates in the middle part of the expectation distribution are not statistically significant. This part of the distribution includes those who have inflation expectation around 2 percent (denoted by the vertical line), which implies that those who have expectations already aligned with the ECB target change them very little when their trust in the ECB increases.

Taken together, results from QR point to the role that trust in the ECB has in anchoring consumers' inflation expectations around the ECB's inflation target. In addition, the fact that estimated effects are stronger among those with high inflation expectations suggests that higher trust in the ECB can be more effective in lowering high inflation expectations than raising low inflation expectations. This finding may be a cause for concern in circumstances like those of the recent past, in which inflation in the euro area was well below the ECB's medium-term target.

We also use an IV approach to the QR estimates, using the same two instruments employed earlier in standard IV. Available IV methods for QR require some modifications of the original specification. In particular, both the trust in the ECB variable and our two instruments have to be redefined as binary variables. Hence, the modified trust in the ECB variable is equal to 0 for values of trust below 7, while it is equal to 1 for values higher than or equal to 7. This implies that a change from 0 to 1 in the modified trust in the ECB variable corresponds to a change of about five units (or 2.3 standard deviations) in the original variable.

The IV QR estimates for various percentiles are shown in figure 9, and also in online appendix table A.2. The estimated coefficients are larger in absolute terms compared with the respective ones from standard QR, as they reflect a much larger underlying change in the trust in the ECB variable. Moreover, IV QR estimates suggest a similar pattern for the role of trust in the ECB across percentiles of the expected inflation distribution to the one derived using standard QR estimates: the results at the upper part of the expectation distribution are quite stronger than those at the lower part, which are statistically significant at percentiles below the 10th one.

Figure 9. IV Quantile Regression Estimates for Expected Inflation



5. Robustness Checks

In this section we examine whether our baseline results pick up the effect of some alternative factors that may influence inflation expectations and its uncertainty.[15] Moreover, we report a few additional robustness checks.

One potential channel through which trust in the ECB can influence inflation expectations is through the knowledge about the ECB's objectives. Earlier research has found a positive correlation between trust in the ECB and knowledge about the ECB (Ehrmann, Soudan, and Stracca 2013). More broadly, general economic knowledge is likely to influence economic expectations (Christensen, van Els, and van Rooij 2006). To that effect, we investigate whether specific knowledge of the ECB's objectives in particular, or economic literacy in general, influences inflation expectations as well as the estimated effect of trust in the ECB.

[15]Detailed results are available from the authors upon request.

To measure knowledge about the ECB's objectives, we ask respondents six true–false questions about these objectives and construct an index representing the number of correct responses (see section A.1 of the online appendix for the wording of these questions).[16] A correlation analysis shows that knowledge and trust are positively associated. Moreover, descriptive statistics results suggest that knowledge on the ECB goals is not broad based, in line with the findings of van der Cruijsen, Jansen, and de Haan (2015). In fact, in most of the questions, around 40 percent of respondents report that they do not know the answer. The average number of correct answers is 3.04, out of a maximum score of 6. Nevertheless, 42 percent of respondents correctly indicate that that an inflation rate below, but close to, 2 percent is a goal of the ECB.[17]

Having an individual measure of institutional knowledge about the ECB allows us to assess whether reported trust in the ECB mainly reflects perceived credibility of the institution or simply knowledge about its role. Before discussing results from formal regression analysis, we note that if responses to the trust in the ECB question reflected credibility rather than knowledge, respondents would report similar trust toward the DNB. Being part of the euro system, the DNB is likely to have similar credibility as the ECB, but obviously it does not have a national inflation mandate anymore. Indeed, responses to the trust in the ECB question are highly correlated with those on a similar question we have asked about trust in the DNB (the correlation is 0.85, and significant at the 1 percent level), which suggests that respondents answer the relevant questions having institutional credibility in mind.

To assess how institutional knowledge affects our results, we add the index of the number of correct questions as an additional covariate to the baseline specification. OLS and IV results, both for expected inflation and variance of inflation expectations, are virtually unchanged. Thus, while knowledge and trust are positively correlated, our results suggest that the institutional credibility aspect

[16] Possible answers include a "do not know" option, which we consider to be equivalent to an erroneous response.

[17] This high number of correct responses may also reflect the fact that the ECB has set an explicit numerical inflation target.

of the trust in the ECB affects inflation expectations over and above knowledge about the ECB goals.

We further examine the role of knowledge by using information on whether respondents answer correctly the questions on the ECB's numerical inflation target. In particular, we distinguish between two groups of respondents. The "has a clue" group consists of those answering correctly the ECB 2 percent inflation target question as well as at least three other questions (out of the six in total). The "no clue" group consists of those not knowing about the 2 percent inflation target and failing to answer correctly three or more questions.[18] When we reestimate our baseline models (i.e., those shown in table 2 and QR), we do not find significant differences in the implied effect of trust in the ECB on inflation expectations between the two groups.

These results are in line with our earlier findings of an independent influence of the institutional credibility on inflation expectations over and above knowledge about the ECB's objectives. With respect to the group that knows about the ECB's numerical inflation target, our findings imply that trusting the central bank can anchor their inflation expectations around it, as they most likely view deviations from the target as temporary ones. As regards the group that does not know about the central bank numerical inflation target, QR results again suggest some anchoring at a broader range around 2 percent, possibly because inflation at this range is compatible with the notion of price stability that this group has. Indeed, responses from a separate survey show that three out of four respondents think that it is most favorable for the Dutch economy if in each year the general level of prices remains stable or increases slightly.

Second, we measure respondents' financial literacy as regards some basic economic concepts using three standard questions extensively used in the related literature (see Lusardi and Mitchell 2014). When including financial literacy in our specifications, the estimated effects of trust in the ECB on inflation expectations are once again unchanged.

[18]Note that the additional restriction of answering at least three other knowledge questions removes from the "has a clue" group clueless respondents whose correct answer that the ECB aims at a price level below but close to 2 percent is due to a guess.

A third factor that may correlate with trust in the ECB and/or inflation expectations is individual optimism. To this end, we first construct an indicator of optimism by taking the difference between subjective life expectancy (i.e., self-reported probability to survive upon age 65, 80, or 90, depending on the current age) and objective life expectancy (by gender and age) in official mortality tables (see also Puri and Robinson 2007, who utilize a similar measure of optimism). In all cases the coefficient of trust in the ECB is hardly affected.[19] Furthermore, we add an alternative measure of optimism in our specification using information from a self-reported measure of optimism.[20] Once more, the estimated effect of trust in the ECB remains unaffected.

Overall, the estimates suggest that trust in the ECB has an effect on economic expectations even controlling for the effect of knowledge about the ECB's functions, financial literacy, and optimism.

Trust in the ECB, and more precisely the institutional credibility or social capital component of trust, may reflect beliefs and values shared by individuals living in the same area. Therefore, we examine whether our baseline results are likely affected by unobserved regional heterogeneity, including possible regional differences in trust in the ECB. To this end, we reestimate our baseline models after accounting for regional fixed effects at the level of the 12 Dutch provinces (which represent the most detailed regional classification available in our data). Results remain virtually unchanged. In addition, we reestimate these models by clustering the standard errors at the province level to allow for possible within-province error dependence. Clustering at such an aggregate level leaves our baseline estimates on trust in the ECB (shown in table 2) statistically significant at 5 percent.

As already mentioned, our estimation sample is smaller than the original one, mainly due to missing observations in inflation and/or

[19]Estimated effects of the trust in the ECB variable remain statistically significant, but the p-value is higher compared with the baseline specification. This is due to the lower number of observations used, as there are missing values of the variable denoting subjective life expectancy.

[20]Respondents indicate to which extent they agree (on a scale from 1 to 5) with the following statement: "Overall, I expect more good things to happen to me than bad things." Guiso, Sapienza, and Zingales (2008) use the same measure to disentangle the effect of trust on stock investing from that of optimism.

trust in the ECB questions. To investigate whether our results are affected by these missing observations, we impute missing values for these two key variables. Adding imputed values increases our estimation sample to 4,248 observations (from 3,055 observations in the baseline estimation using the simple triangular distribution). Results from this larger sample, based on OLS and QR estimates of the relationship between trust in the ECB and inflation expectations, are similar to those we present in section 4 (results on expected inflation and its variance are reported in online appendix table A.3).

As a final robustness check, we estimate the baseline specification replacing the simple triangular distribution with the split one. We find that the results on inflation expectations, uncertainty, and anchoring remain unchanged (see online appendix table A.4 for these results). Similarly, results are robust to assuming alternative distributional forms, such as uniform and split uniform distributions.

6. GDP Growth

The broader objectives of monetary policy are not limited to price stability but include also other macroeconomic variables, among which economic growth is of course prominent. Hence, it is instructive to examine whether trust in the ECB affects also individual expectations regarding economic growth and economic fluctuations (e.g., GDP growth rate volatility). To elicit the distribution of expected GDP growth, we ask individuals the same sequence of three questions as for inflation: the minimum and maximum expected growth rate, and the chance that growth exceeds the midpoint of the reported minimum and maximum.

As is the case with inflation expectations, we plot the mean of expected growth and its variance by bins of trust in the ECB in figures 10 and 11, respectively. There is a clear positive association between trust in the ECB and individual expectations about economic growth. On the other hand, we find essentially no association between trust in the ECB and the variance of expected growth.

The patterns shown in the two figures are confirmed when we estimate multivariate regressions with expected growth and its variance as dependent variables. Table 3 reports OLS and IV regressions for the mean (columns 1–4) and variance (columns 5–8) of expected

Figure 10. Average Expected GDP Growth Rate by Levels of Trust in the ECB

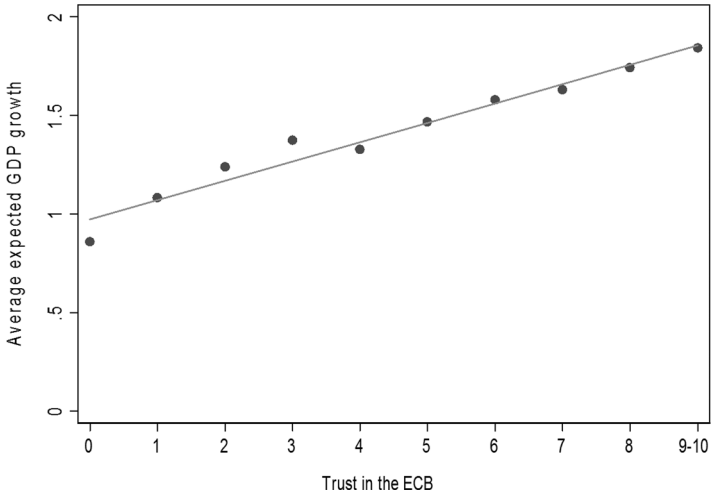
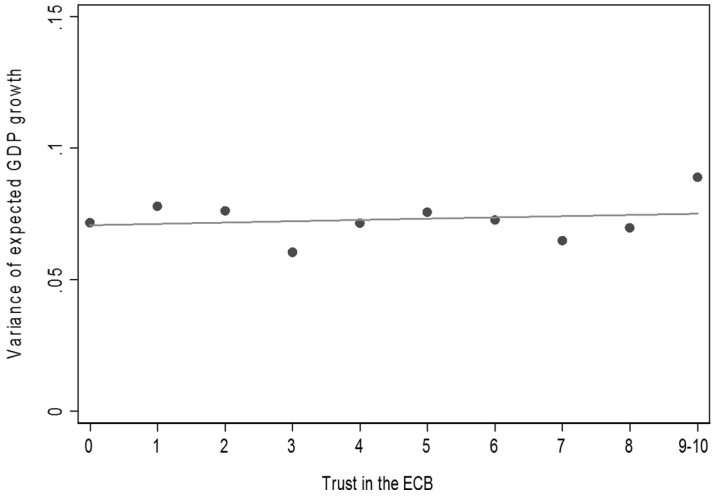


Figure 11. Variance of Expected GDP Growth Rate by Levels of Trust in the ECB



GDP growth. As is the case with inflation expectations, we cluster Huber-White robust standard errors at the household level, and winsorize the dependent variables at the top and bottom 1 percent

Table 3. OLS and IV Results for Expected Growth and Its Variance

	Mean of Expected Growth Rate				Variance of Expected Growth Rate			
	OLS		IV		OLS		IV	
Variable	**Coeff. (1)**	**Std. Error (2)**	**Coeff. (3)**	**Std. Error (4)**	**Coeff. (5)**	**Std. Error (6)**	**Coeff. (7)**	**Std. Error (8)**
Trust in the ECB	0.0792	0.0106***	0.1250	0.0347***	−0.0010	0.0010	−0.0030	0.0050
Age	−0.0104	0.0102	−0.0088	0.0107	−0.0040	0.0020**	−0.0030	0.0020*
Age Squared	0.0001	0.0001	0.0001	0.0001	0.0000	0.0000	0.0000	0.0000
Female	−0.1481	0.0433***	−0.1950	0.0466***	0.0130	0.0060**	0.0100	0.0060
Couple	0.0336	0.0682	0.0415	0.0707	0.0080	0.0090	−0.0010	0.0100
Household Size	−0.0034	0.0283	−0.0081	0.0295	−0.0030	0.0040	0.0000	0.0040
High School Graduate	0.0520	0.0631	0.0786	0.0692	−0.0040	0.0080	0.0010	0.0090
College Graduate	0.0738	0.0595	0.0710	0.0745	0.0040	0.0080	0.0120	0.0100
Logarithm of Household Net Income	0.0220	0.0232	0.0287	0.0243	0.0020	0.0030	0.0040	0.0020**
Constant	0.9066	0.3309***	0.5269	0.3845	0.2060	0.0540***	0.1780	0.0600***
Region/Wave Dummies	Yes		Yes		Yes		Yes	
First-Stage F-Test			92.332				92.332	
Endogeneity Test			2.070				0.019	
P-value			0.150				0.890	
Test of Overidentifying Restrictions			0.297				1.419	
P-value			0.586				0.234	
Observations	3,145		2,716		3,145		2,716	

Notes: This table shows OLS and IV estimation results from models using the expected growth rate and its variance (both measured in percentage points) as dependent variables. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively.

of observations.[21] The test statistics for the IV regressions suggest again that the instruments are strong and that one cannot reject instrument exogeneity. At the same time, the Hausman test does not reject the null hypothesis that trust in the ECB is exogenous. Therefore, one could restrict attention to the OLS estimates. Yet, to maintain consistency with our analysis of expected inflation and its variance, we present both OLS and IV results. In any case, both analyses lead to qualitatively similar estimates of the effects of trust in the ECB.

In particular, we find that trust in the ECB increases expected GDP growth. According to the baseline OLS estimate, a one-standard-deviation higher trust in the ECB implies an increase in expected GDP growth by 17 basis points. The coefficient is statistically significant (at the 1 percent level) and the effect corresponds to 12 percent of the sample mean of expected GDP growth. The corresponding IV estimate is somewhat higher (27 basis points, corresponding to 18 percent of the sample mean). In contrast, results in columns 5–8 imply that trust has no effect on the uncertainty about expected GDP growth.[22] As regards other covariates, we find a negative (positive) association between the female dummy and expecting a higher (being more uncertain about the) growth rate. This seems aligned with insights from the cognitive psychology literature on the link between gender and reactions related to anxiety about the future (e.g., Robichaud, Dugas, and Conway 2003).

The finding that trust in the ECB affects inflation expectations negatively at the mean, and growth expectations positively, suggests that respondents on average do not associate necessarily higher inflation with higher growth. This could be due to several reasons, including the "Great Moderation" period that featured both low inflation and economic prosperity, memories of past stagflation, or

[21] The number of observations is slightly higher, as the number of "do not know" responses to the GDP growth questions is lower than for the inflation questions (see also Christensen, van Els, and van Rooij 2006).

[22] We have also estimated a series of QR to examine whether estimates of trust in the ECB display a pattern of anchoring around a specific value of the expected GDP growth distribution. We do not find any evidence for anchoring in this case. QR estimates are always positive and decline across quantiles, suggesting that higher trust in the ECB is associated with higher expected GDP growth, and more so among those with low GDP growth expectations.

experience drawn from other countries, where deep recessions have been accompanied by episodes of high inflation rates.

7. Summary

Historically, central banks have paid a lot of attention to inflation expectations formed in financial markets and by professional forecasters. More recently, central banks have shown increased interest in consumer expectations and beliefs because they can help make policy more effective (Blinder et al. 2008; Bernanke 2013). In this paper, we investigate the extent to which trust in the ECB affects individuals' expectations and uncertainty about future inflation and induces inflation anchoring at the ECB's inflation target of below, but close to, 2 percent. The empirical evidence draws upon a special questionnaire module introduced in a recent survey of a representative sample of the Dutch population.

A first finding of our analysis is that a high level of trust lowers inflation expectations. This result may be due to the fact that traditionally central banks have been mainly concerned about inflation exceeding their target and communicated to the public their commitment to raise interest rates to restrain inflation. Consequently, a high level of public trust in the ECB is likely to reflect trust in the ECB's commitment and ability to fight high inflation and thereby induces lower inflation expectations on average.

Recently, central banks have been using unconventional policy instruments to cope with a long period of low inflation and near-zero interest rates. Based upon a survey among central bank presidents and academic experts, Blinder et al. (2017) argue that these new policy instruments as well as the increased use of communication will permanently remain in the toolkit of central banks. Communication is, among others, important for central banks for anchoring inflation expectations around the target inflation rate and preventing medium-term inflation expectations from falling below target.

Our findings are directly related to this desired anchoring of inflation expectations because they indicate that trust in the ECB induces anchoring around the medium-term inflation target. Specifically, we show that the effect of trust is not uniform across the distribution of inflation expectations: at the lower end of the distribution,

an increase in trust increases inflation expectations, while the opposite is true at the higher end of the distribution of inflation expectations. Estimated effects are particularly strong among respondents having high inflation expectations, suggesting that higher trust in the ECB matters more for lowering high inflation expectations as opposed to increasing low inflation expectations. This finding may be a cause for concern in a low interest rate environment, especially as regards the segment of the population with quite low inflation expectations. It suggests that central banks may benefit from effectively communicating their commitment to raise prices toward their target as forcefully as their commitment to fight high inflation in previous times.

In a related vein, we also find that trust in the ECB reduces individual uncertainty about future inflation, thus contributing to public confidence about future price stability and the economy's prospects.[23] Taken together, our findings suggest that a high level of trust supports the monetary policy of the ECB because it contributes to the anchoring of inflation expectations in the population around the target of below, but close to, 2 percent.

One may argue that it is not the institutional credibility of the ECB that matters for our findings, but rather the knowledge about the tasks and the goals of the ECB (or the knowledge about economic concepts in general). While there is a positive association between knowledge of the ECB goals and trust in the ECB, our results are virtually unaffected when knowledge about the ECB's objectives and financial literacy are taken into account. Hence, it appears that the institutional credibility component of trust in the ECB has an independent influence on inflation expectations.

Finally, our findings suggest that the effectiveness of monetary policy could benefit from investing in the buildup of trust and institutional credibility. In the current environment of low interest rates, where standard monetary policy measures are difficult to implement

[23] From a monetary policy perspective, reduced uncertainty strengthens anchoring around a medium-term inflation target and induces equilibrium prices to converge faster toward this target (Bernanke 2013). From a broader perspective, reduced uncertainty is beneficial for economic welfare. It helps households' financial planning and lowers the need for precautionary saving (see Christelis et al. 2020 on the effect of consumption uncertainty on the latter).

and likely to be less effective, such an investment is likely to be particularly beneficial.

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What Rule for the Federal Reserve? Forecast Targeting*

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Forecast targeting means selecting a policy rate and policy rate path so that the forecasts of inflation and employment "look good," that is, best fulfill the Federal Reserve's dual mandate. It means publishing, explaining, and justifying the path and forecasts. This contributes to their credibility and to effective policy implementation. External observers can review policy and hold the Federal Reserve accountable, both in real time and after the fact. Forecast targeting uses all relevant information and allows adaption to new circumstances and therefore performs better than a Taylor-type rule. The Federal Reserve is arguably to a considerable extent already practicing forecast targeting. Forecast targeting can be used with the new monetary policy strategy of flexible average inflation targeting announced in August 2020.

JEL Codes: E52, E58.

"*The Fed has a rule.* The Fed's rule is that we will go for a 2% inflation rate; we will go for the natural rate of unemployment; we put equal weight on those two things; we will give you information about our projections, our interest rate. *That is a rule*, and that is a framework that should clarify exactly what the Fed is doing." (Bernanke 2015b)

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1. Introduction

How should the Federal Reserve conduct monetary policy so as to best fulfill its mandate of price stability and maximum employment? What decisionmaking process should the Federal Reserve follow, what information should it rely on, and how should it set its policy instruments? What of its information, deliberations, and decision should the Federal Reserve publish? How can the Federal Reserve's policy conduct best be reviewed and how can the Federal Reserve most effectively be held accountable for fulfilling its mandate?

These issues are arguably always of importance, but they became more urgent in the context of proposed legislation in 2015 and 2017 by the U.S. Congress. According to the Fed Oversight Reform and Modernization (FORM) Act (U.S. Congress 2015) and, with identical words, the Financial CHOICE Act (U.S. Congress 2017), the Federal Open Market Committee (FOMC) Chair shall within 48 hours after each FOMC meeting submit a "Directive Policy Rule" (DPR) which identifies the "Policy Instrument" and "includes the coefficients" through which the "Intermediate Policy Inputs" determine the level of the policy instrument. In particular, the DPR shall

> include a statement as to whether the Directive Policy Rule substantially conforms to the Reference Policy Rule and, if applicable, (A) an explanation of the extent to which it departs from the Reference Policy Rule; (B) a detailed justification for that departure; ... (U.S. Congress 2015, section 2).

Importantly, the Reference Policy Rule is specified in words and numbers to be the standard Taylor (1993) rule for the federal funds rate,

$$i_t = 2 + \pi_t + 0.5\,(\pi_t - 2) + 0.5\,y_t, \tag{1}$$

where i_t denotes the federal funds rate in quarter t, π_t denotes inflation over the previous four quarters, and y_t denotes the gap between gross domestic product (GDP) and an estimate of potential GDP.

Clearly, these provisions in the legislation make the Taylor rule the benchmark for the Federal Reserve's monetary policy, and if there are any departures from the rule, these departures require "a detailed justification." Furthermore, the Government Accountability Office (GAO) would be responsible for determining whether the

FOMC's DPR would meet all the legislation's criteria. Any time the FOMC's DPR was judged not to be in compliance with the GAO-approved rule, or any time the FOMC just changed its DPR, the GAO would have to conduct a full review of monetary policy and submit a report to Congress.

As explained in a letter from Chair Yellen to Congress (Yellen 2015), for several reasons the provisions of the FORM Act would severely impair the Federal Reserve's ability to carry out its congressional mandate to promote effectively the goals of maximum employment and stable prices. One obvious reason is that the provisions would threaten the Federal Reserve's considerable independence in deciding how to best fulfill its mandate. The provisions would effectively put Congress and the GAO in the role of reviewing short-run policy decisions and in a position to influence those decisions in real time. There is considerable theoretical and ample historical evidence that such short-run political interference in monetary policy leads to poor economic outcomes.

Another reason is that there are considerable problems with Taylor-type rules that make them lead to poor economic outcomes in many situations. A Taylor-type rule is too restrictive and mechanical, does not take into account all relevant information, and lacks the flexibility required to handle complex and changing situations.

More precisely, as discussed in Svensson (2003b), first, a Taylor-type rule is not optimal, in the sense of best stabilizing both inflation around the inflation target and unemployment around its long-run sustainable rate, and in some circumstances it is far from optimal. A Taylor-type rule makes the policy rate respond with some fixed coefficients to the current inflation gap and either the current GDP gap or the current unemployment gap.

But good monetary policy needs to respond to much more information than is contained in the current observations of those gaps. In particular, in order to best fulfill the mandate of maximum employment and price stability, it is not sufficient for the policy rate to respond only to the current levels of inflation, GDP, and employment or unemployment. Instead, optimal policy requires a response to the *determinants* of the future realizations of inflation and employment. These determinants normally include the current levels of inflation, GDP, and employment or unemployment but, importantly, also many other variables and shocks. Second, the relevant

information depends on circumstances and changes over time. In order to best achieve the mandate, there is a crucial role of *judgment* (information, knowledge, and views outside the scope of a particular model) in modern monetary policy, and the appropriate use of good judgment can dramatically improve monetary policy performance (Svensson 2005). But a Taylor-type rule leaves no room for judgmental adjustments. Third, the beneficial development of monetary policy due to learning and new information will conflict with the legislation of (or commitment to) a particular Taylor-type rule. Finally, in spite of considerable academic work and promotion, no central bank has actually chosen to commit itself to a Taylor-type rule (and prominent central bankers are very critical of the idea). In short, a Taylor-type rule is not optimal, in the sense of best achieving the mandate, and is too rigid to adapt to changing circumstances.[1]

A possible answer to these problems, in particular to the second one mentioned above, is that a Taylor-type rule should not be followed mechanically. Instead, deviations from the rule are allowed. The rule should be seen as mere "guidelines" for monetary policy. This is the view expressed in the original proposal of Taylor (1993) and, in more detail, in Taylor (2000). A problem with this answer is that the rule is then *incomplete*: some deviations are allowed, but *there are no rules for when deviations from the Taylor-type rule are appropriate*. As discussed in some detail in Svensson (2003b), this arguably makes the use of simple instrument rules such as Taylor-type rules as mere guidelines for monetary policy too vague to be operational.[2]

What rule for the FOMC's monetary policy setting would then better fulfill the Federal Reserve's mandate over time and also make

[1]Federal Reserve Board (2019) discusses the Federal Reserve's views on and current use of different policy rules. The section "Systematic Monetary Policy in Practice" provides similar criticism of Taylor-type rules as here.

[2]Walsh (2015, 2017) provides a discussion of the pros and cons of goals-based and rules-based monetary policy (the latter there meaning commitment to a Taylor-type rule) in a New Keynesian model in which the central bank is subject to political pressure or for other reasons deviates from the socially optimal goals. In particular, he shows that equilibrium central bank deviations from the socially optimal policy or from the Taylor-type rule then depend on these pressures and the degree of commitment to the Taylor-type rule. He refers to this as a "rule" for the deviations from the Taylor-type rule. I find it difficult to see this specific result as a general operational and verifiable rule for such deviations.

the Federal Reserve's policy sufficiently transparent so that the Federal Reserve can be held accountable for fulfilling the mandate? This paper suggests that *forecast targeting* is likely to allow the FOMC to effectively fulfill its mandate as well as to be held accountable for fulfilling the mandate. Indeed, forecast targeting is the rule that Bernanke (2015b) briefly refers to in the quote at the beginning of this paper, suggesting that the Federal Reserve is already to some extent practicing forecast targeting.[3] Forecast targeting means setting the policy rate (the federal funds rate) and the policy rate path so that the resulting forecasts for the Federal Reserve's "target variables"—inflation and employment (or unemployment)—best fulfill the Federal Reserve's mandate of maximum employment and price stability. Forecast targeting also involves publishing and justifying the FOMC's policy rate path and forecasts for inflation and employment. This serves to effectively implement the selected policy in order to make it credible with the financial market and other economic agents, as well as to make it possible to hold the Federal Reserve accountable for fulfilling its mandate.[4]

This paper shows how to apply forecast targeting to the Federal Reserve's specification of its previous monetary policy strategy in FOMC (2019b). It was written before the Federal Reserve's announcement in August 2020 of a new strategy of "flexible average inflation targeting," where the price-stability mandate was modified to achieve "inflation that averages 2 percent over time" and the maximum-employment mandate was modified to mitigate only "shortfalls" of employment from its maximum level, not "deviations" as previously (FOMC 2020; Powell 2020). As argued in Svensson (2020), forecast targeting is equally suitable to the new strategy of flexible average inflation targeting. A new section, section 6, has been added to this paper that explains how forecast targeting can be modified to be suitable for the new strategy.

To clarify how forecast targeting works, consider for simplicity a situation of relatively normal times when the Federal Reserve is

[3]Forecast targeting rather than Taylor-type rules is discussed and promoted in, for example, Svensson (1997, 2003b, 2011), Bernanke (2004, 2015a), Woodford (2004, 2007, 2012), Qvigstad (2005), and Kohn (2007, 2012). The term "inflation-forecast targeting" was introduced in Svensson (1997), and the term "forecast targeting" in Svensson (2003b).

[4]Rudebusch and Williams (2008) provide early support for publishing the Federal Reserve's policy rate projection.

not doing any active balance sheet policy but is only using a policy (interest) rate, the federal funds rate, as its policy instrument. Furthermore, assume for simplicity that the labor market participation rate is independent of monetary policy, so that for monetary policy purposes employment varies negatively one-to-one with unemployment. Then maximum employment corresponds to what FOMC (2019b) calls the longer-run normal rate of unemployment, what I will call the (minimum) long-run sustainable unemployment rate. Under this simplification, the Federal Reserve's mandate is to keep inflation close to its target of 2 percent and unemployment close to its estimated long-run sustainable unemployment rate.

Two important circumstances then need to be taken into account: First, monetary policy actions tend to influence economic activity and prices with a lag. Therefore, monetary policy is more effective in fulfilling the mandate if it is guided by *forecasts of future* inflation and unemployment rather than by current inflation and unemployment.

Second, the current policy rate has a very small direct impact on economic activity and prices. What matters for economic activity and prices is instead market *expectations of future* policy rates. These expectations affect longer-term interest rates and asset prices, which in turn have an impact on activity and prices. It is the entire expected path of future policy rates that affects economic activity, not the policy rate over the next few days and weeks. This means that an effective monetary policy decision cannot only consist of setting the current policy rate; it must explicitly or implicitly also involve the selection of a policy rate path, a forecast of the future policy rate. Not to discuss and select a policy rate path is an incomplete decisionmaking process (Svensson 2007a).

Given this, a rule for the FOMC that effectively fulfills its mandate is to select a policy rate and a policy rate path so that the resulting forecasts for inflation and unemployment "look good." Here, "looking good" means best fulfilling the Federal Reserve's mandate, that is, best stabilizing inflation around its target and unemployment around its long-run sustainable rate.[5]

[5]Some technical issues to ensure determinacy and the terminal conditions on the chosen policy rate path are discussed in section 3.5.

Why is this rule, forecast targeting, better than a Taylor-type rule? First, it takes into account all relevant information available to the Federal Reserve. It takes into account the information about the economy, economic activity, and prices that has an impact on the forecasts of inflation and unemployment at a given policy rate path. It also takes into account all relevant information about the transmission mechanism of monetary policy, that is, how changes in the policy rate path affect the forecasts of inflation and unemployment at given information about the current state of the economy. Second, the rule therefore adapts to new information and changes in circumstances, and it allows for judgmental adjustments. It avoids the restrictiveness and inflexibility of a Taylor-type rule. The selected policy rate path and forecasts of inflation and unemployment will in practice be a combination of model simulations, sometimes from several models, and judgmental adjustments.

However, for successful implementation and realization of the selected policy, the policy rate path needs to be credible, in the sense of market expectations of future policy rates being aligned with the policy rate path. Implementation of monetary policy is largely about the management of expectations (Woodford 2004). This includes making the *actual* financial conditions align with the *intended* financial conditions, where the latter can be seen as represented by the policy rate path. Economic agents' expectations of future inflation also matter. If the FOMC manages to make the inflation target credible, in the sense of making economic agents' inflation expectations align with the inflation target, stabilization of inflation around its target is easier, because actual inflation is much affected by previous expectations of inflation. Then it is also easier to stabilize unemployment around its long-run sustainable rate. The tradeoff between stability of inflation around the target and of unemployment around its long-run sustainable rate becomes more favorable.

The most effective contribution to making the policy rate path credible with the market participants and other economic agents is to publish the policy rate path and the forecasts of inflation and unemployment and justify them and the policy decision. Not to publish the policy rate path would be to hide the most important information (Svensson 2007a). Forward guidance is then the default, in the sense that there is forward guidance in the form of a published policy

rate path. Normally, this is a forecast conditional on current information, not a commitment. Therefore, like other forecasts, it often changes at the next decision. In exceptional situations, for example, when the Federal Reserve is restricted by the effective lower bound (ELB) for the policy rate, it may be a commitment through a certain date (time dependent) or conditional on a specific outcome of inflation or employment (state dependent).[6]

A common argument against publishing a policy rate path is that market participants would not understand that it is a conditional forecast that changes over time but interpret it as commitment and be confused when it changes. The Reserve Bank of New Zealand, the Bank of Norway, and the Riksbank have published their policy rate paths since 1997, 2005, and 2007, respectively. There is thus considerable experience of such publication. I am not aware that market participants in these countries have had any difficulties understanding that the policy rate path is a forecast, not a commitment. The Riksbank used to repeat at every publication of the path that "it is a forecast, not a promise." After a few years, these reminders were certainly redundant. The Federal Reserve could repeat the same phrase, if there would be concerns about misunderstandings.

In addition to justifying how new information since the last decision has affected the forecasts and the selected policy rate path, the justification of the decision may include demonstration of why the inflation and unemployment forecasts "look good," that is, best fulfill the Federal Reserve's mandate. If required, this can be done by showing that other policy rate paths than the one selected lead to inflation and unemployment forecasts that look less good, that is, do not fulfill the mandate as well. This can be done more explicitly with the use of what can be called mean squared gaps, which quantify the deviation of the inflation forecast from the inflation target and the deviation of the unemployment forecast from the long-run sustainable unemployment rate (Svensson 2011).

It is common to argue that central banks should convey their reaction function to the market participants and other economic agents. However, under forecast targeting the reaction function, meaning how the policy rate and the policy rate path respond to

[6]See Bernanke (2017) for further discussion.

information available to the central bank, is far too complex to write as a simple formula such as a Taylor-type rule. It is actually too complex to write down, period—that is, as a mathematical function describing precisely how the policy rate responds to different kinds of information. The policy rate and policy rate path will normally respond to all relevant information, that is, all information that shifts the forecasts of inflation and unemployment. This is a long and changing list, with response coefficients that cannot be specified in advance.

But the reaction function can be conveyed in more general and approximate but still both systematic and simple terms. If initially the forecasts look good, for any piece of information that shifts the inflation forecast up (down) and/or shifts the unemployment forecast down (up), policy will normally be tightened (eased), meaning that the policy rate path will shift up (down). If this response is understood by and credible with the market participants, any new information that is deemed to shift up (down) the inflation outlook or shift down (up) the unemployment outlook may result in a market response that shifts up (down) the yield curve. This way the financial conditions may shift in the appropriate direction and perhaps even of the appropriate magnitude even before the central bank has responded with a new policy rate and policy rate paths at the next decision.[7]

Finally, the publication and justification of the FOMC's policy rate path and inflation and unemployment forecasts make it possible to hold the FOMC accountable for fulfilling the mandate. The policy rate path and forecasts of inflation and unemployment, the FOMC's justification of them, and its fulfillment of its mandate can be scrutinized and reviewed both in real time and after the fact, that is, after the outcome for inflation and unemployment have been observed, by external observers and experts and at the usual hearings in congressional committees (Svensson 2012).

Altogether, forecast targeting can be seen as a case of "constrained discretion" (Bernanke and Mishkin 1997), where the constraint to fulfill the mandate is made most explicit.

[7]Federal Reserve Board (2019, p. 39) makes a very similar point.

In the rest of the paper, section 2 discusses the interpretation and specification of the Federal Reserve's mandate. Section 3 discusses how the Federal Reserve can effectively fulfill its mandate by the decisionmaking process of forecast targeting. Section 3.2 provides an example with a comparison of the performance of forecast targeting and a Taylor rule when there are expectations of future shocks. Section 3.3 discusses the implementation of the policy decision, and section 3.4 summarizes the forecast-targeting policy rule in three steps. Sections 3.5 and 3.6 discuss how the issues of determinacy and time consistency can be handled. Section 3.7 clarifies the reaction function under forecast targeting, and section 3.8 clarifies the distinction between forecast targeting, instrument rules, and targeting rules. Section 4 discusses how the Federal Reserve can be held accountable. Section 4.1 shows an example, from the Riksbank's monetary policy decision in February 2013 (Sveriges Riksbank 2013), of how alternative policy rate paths and corresponding forecasts of inflation and unemployment (including mean squared gaps) can be used to examine whether the decision best fulfills the mandate. Section 5 includes a discussion of to what extent the Federal Reserve is already practicing forecast targeting. Section 6 shows how forecast targeting can be modified to be suitable for the Federal Reserve's new monetary policy strategy of flexible average inflation targeting. Section 7 provides some conclusions.

2. The Mandate

The one-page well-written FOMC "Statement on Longer-Run Goals and Monetary Policy Strategy" (FOMC 2019b, first adopted in January 2012) clarifies the Federal Reserve's monetary policy goals and strategy. The Federal Reserve's statutory mandate is to promote maximum employment and price stability.[8] The FOMC has decided that a symmetric 2 percent inflation target is most consistent over the longer run with its statutory mandate. Regarding maximum

[8]More precisely, Congress has given the Federal Reserve the statutory mandate "to promote effectively maximum employment, stable prices, and moderate long-term interest rates." Moderate long-term interest rates will normally follow from low and stable inflation.

employment, the FOMC notes that the maximum level of employment, in contrast to the rate of inflation, is largely determined not by monetary policy but by nonmonetary factors that affect the structure and dynamics of the labor market. These factors may change over time and may not be directly measurable. Consequently, it would not be appropriate to specify a fixed goal for employment; rather, the maximum level of employment must be estimated from a range of indicators, and such estimates are uncertain and subject to revision. An important indicator is the FOMC's estimate of what it calls the longer-run normal rate of unemployment.

The FOMC provides further information on how it sets monetary policy:

> In setting monetary policy, the Committee seeks to mitigate deviations of inflation from its longer-run goal and deviations of employment from the Committee's assessments of its maximum level. These objectives are generally complementary. However, under circumstances in which the Committee judges that the objectives are not complementary, it follows a balanced approach in promoting them, taking into account the magnitude of the deviations and the potentially different time horizons over which employment and inflation are projected to return to levels judged consistent with its mandate. (FOMC 2019b)

Given this, the mandate can be well formalized by a standard quadratic loss function of inflation and employment. If, for simplicity, the labor market participation rate is assumed to be independent of monetary policy, maximum employment can be replaced by the (minimum) longer-run normal unemployment rate. The mandate can then be expressed in terms of a standard quadratic loss function of inflation and unemployment. Furthermore, a "balanced approach," and the explicit statement of former Chair Bernanke (2015b) in the quote at the beginning of this paper, can be interpreted as an equal weight on stabilization of inflation and on stabilization of unemployment. Indeed, then Vice Chair Yellen (2012, p. 13) stated:

> The balanced-approach strategy endorsed by the FOMC is consistent with the view that maximum employment and price

> stability stand on an equal footing as objectives of monetary policy.

Furthermore, Vice Chair Clarida (2019, p. 5) notes:

> As a practical matter, our current strategy shares many elements with the policy framework known in the research literature as "flexible inflation targeting." However, the Fed's mandate is much more explicit about the role of employment than that of most flexible inflation-targeting central banks, and our statement reflects this by stating that when the two sides of the mandate are in conflict, neither one takes precedent over the other.

Then the quarter-t loss, L_t, can be represented by the quadratic loss function,

$$L_t = (\pi_t - \pi^*)^2 + (u_t - u^*)^2, \tag{2}$$

where π_t denotes the inflation rate, π^* denotes the 2 percent inflation target, u_t denotes the unemployment rate, and u^* denotes the FOMC's (latest) estimate of the longer-run normal unemployment rate, which I will call the (minimum) long-run sustainable unemployment rate. Furthermore, the inflation rate, π_t, and the unemployment rate, u_t, can be seen as the two target variables of monetary policy (target variables are the variables that enter the loss function). Here we should not forget the important difference, mentioned above, that the target for inflation, π^*, is determined by the FOMC, but the target for unemployment, the long-run sustainable unemployment rate, u^*, is *estimated*, not determined, by the FOMC, because it is determined largely by nonmonetary structural factors beyond the control of monetary policy.[9]

This quadratic loss function thus implies the independent stabilization of inflation around the inflation target and the unemployment rate around the longer-run normal rate. Furthermore, in line with the quotes above of Yellen (2012), Bernanke (2015b), and Clarida (2019), the two objectives have equal weight in the loss function. Given this, it is arguably a good representation of the dual mandate.

[9]Because the FOMC's estimate of the long-run sustainable rate may change over time, it could be indexed by the quarter of the latest estimate.

A quadratic loss function of inflation and unemployment can also be derived as a quadratic approximation to household welfare in some simple models. But this is less relevant here. The relevant property here is that the loss function is a good representation of the Federal Reserve's legislated mandate.

In a given quarter t, the mandate for the future can then be formalized as setting monetary policy so as to minimize the intertemporal loss function

$$\mathrm{E}_t \sum_{\tau=0}^{T} \delta^\tau L_{t+\tau} = \mathrm{E}_t \sum_{\tau=0}^{T} \delta^\tau [(\pi_{t+\tau} - \pi^*)^2 + (u_{t+\tau} - u^*)^2], \qquad (3)$$

where E_t denotes FOMC expectations conditional on its information in quarter t, T denotes a finite horizon (measured in quarters), and δ is a discount factor that satisfies $0 < \delta \leq 1$ and in practice is very close to or equal to one.[10]

3. Fulfilling the Mandate

What rule for the FOMC's monetary policy setting may then effectively fulfill the Federal Reserve's mandate over time? That is, what rule would minimize the intertemporal loss, (3)? Here, given the problems of Taylor-type rules noted in section 1 (examined in more detail in Svensson 2003b), I consider more general rules, rules in the sense of "prescribed guides for monetary policy."

Let us simplify somewhat by considering a situation of normal times, when the Federal Reserve is not doing any active balance sheet policy but is only using the policy (interest) rate, currently the federal funds rate, as its policy variable. Furthermore, let us, as above, consider inflation and unemployment as the two target variables.[11]

[10]The horizon, T, can in theory be infinite, but in practice it is finite—for example, 20 quarters. Central banks often publish forecasts for up to 12 quarters. A finite horizon also implies that the intertemporal loss function converges not only for $0 < \delta < 1$ but also for $\delta = 1$.

[11]That is, under the assumption that the labor market participation rate is approximately independent of monetary policy. If it is not, the loss function should be expressed in terms of the employment rate instead of the unemployment rate.

Two important circumstances then need to be taken into account. First, monetary policy actions tend to influence economic activity and prices with a lag. Monetary policy has a small or zero impact on inflation and unemployment in the current quarter. The major impact is in future quarters. Therefore monetary policy is more effective in fulfilling its mandate if it is guided by *forecasts of future inflation and unemployment* rather than (estimates of) current inflation and unemployment.[12]

Second, the current policy rate has a very small direct impact on economic activity and prices. What matters for economic activity and prices is instead market expectations of future policy rates. These expectations affect longer-term interest rates and asset prices, which in turn have an impact on activity and prices. It is the entire expected path of future policy rate that affects economic activity, not policy rate over the next few days and weeks.

3.1 The Monetary Policy Decision

It follows that an effective monetary policy decision cannot only consist of setting the current policy rate; it must explicitly or implicitly also involve the selection of an expected path for the future policy rate. Given this, the rule for the FOMC that would best fulfill its mandate is to select a policy rate path such that, conditional on this path and current information about the economy, the resulting forecasts for inflation and unemployment "look good." Here, "looking good" means best fulfilling the Federal Reserve's mandate, that is, best stabilizing inflation around its target and unemployment around its long-run sustainable rate.

Let me make this a bit more precise with some notation and definitions, following Svensson (2011). First, let $i^t \equiv (i_{t,t}, i_{t+1,t}, \ldots, i_{t+T,t}) \equiv \{i_{t+\tau,t}\}_{\tau=0}^{T}$ denote the policy rate path in the current quarter t. Here $i_{t,t}$ denotes the current policy rate and $i_{t+\tau,t}$ for $\tau = 1, 2, \ldots, T$ denotes the FOMC's quarter-t forecast of, or plan for, the policy rate in future quarters $t + \tau$. Second, let

[12]One should remember that current inflation and unemployment are not directly observed in real time. The numbers published by statistics authorities are therefore also estimates, "nowcasts," of the "true" current inflation and unemployment.

$\pi^t \equiv \{\pi_{t+\tau,t}\}_{\tau=0}^T$ and $u^t \equiv \{u_{t+\tau,t}\}_{\tau=0}^T$ denote the FOMC's forecasts of inflation and unemployment.

Under forecast targeting, these forecasts should be (probability) mean forecasts, not modal forecasts. Modal forecasts imply a rather bizarre "perfectionist" loss function, which is difficult to defend (Svensson 2003a, section 6). Note that, relative to a modal forecast, a mean forecast can be seen as a risk-adjusted forecast. Mean forecasts are sufficient for optimal policy if the conditions for so-called certainty equivalence are fulfilled (a linear model, additive shocks, and a quadratic loss function). It is not obvious to what extent mean forecast targeting (relying on certainty equivalence and hence only on mean forecasts) is still an acceptable approximation when there is model uncertainty, multiplicative uncertainty, and so on (Brainard 1967; Söderström 2002). Depending on the precise nature of the uncertainty, optimal policy may be more or less aggressive than the certainty-equivalent policy. My experience is that, in many practical situations, the information available is not sufficient to decide in what direction optimal policy deviates from the certainty-equivalent one, in which cases I find it justified to stay with the certainty-equivalent policy. The main exception is the nonlinearity caused by the ELB on nominal interest rates. When there is a risk that the ELB will bind in the future, policy should—all else being equal—normally be more expansionary than the certainty-equivalent one. The central bank should not "keep its powder dry." The ELB will not prevent the central bank from tightening policy in response to future positive shocks that increase inflation and reduce unemployment. But it may prevent the central bank from easing policy sufficiently in response to future large negative shocks. This results in a downward (upward) bias and volatility of future inflation (unemployment) that—all else being equal—justifies some current policy easing, so as to increase the future distance to the ELB and thereby reduce the risk of a future binding ELB (Reifschneider and Williams 2000; Williams 2019).[13]

Third, define the forecast loss, $L_{t+\tau,t}$, as

$$L_{t+\tau,t} = (\pi_{t+\tau,t} - \pi^*)^2 + (u_{t+\tau,t} - u^*)^2. \tag{4}$$

[13]Clouse (2018) provides a pedagogical discussion of some of these issues.

It represents the loss from deviations of quarter-t forecasts of quarter-$(t+\tau)$ inflation and unemployment from, respectively, the inflation target and the long-run sustainable unemployment rate. Then the quarter-t intertemporal forecast loss, $\mathcal{L}_t$, is given by

$$\mathcal{L}_t = \sum_{\tau=0}^{T} L_{t+\tau,t} = \sum_{\tau=0}^{T} (\pi_{t+\tau,t} - \pi^*)^2 + \sum_{\tau=0}^{T} (u_{t+\tau,t} - u^*)^2, \quad (5)$$

where the discount factor, δ, for simplicity has been set equal to one.

Furthermore, the deviations of the inflation forecast from its target and the unemployment forecast from its long-run sustainable rate can be measured by the mean squared gaps for inflation and unemployment, defined as follows. The intertemporal forecast loss, (5), divided by the horizon, can be written

$$\mathcal{L}_t/T = \mathrm{MSG}_t^{\pi} + \mathrm{MSG}_t^{u}, \quad (6)$$

where MSG_t^{π} and MSG_t^{u} denote the mean squared gaps (MSGs) for, respectively, inflation and unemployment and are defined as

$$\mathrm{MSG}_t^{\pi} \equiv \sum_{\tau=0}^{T} (\pi_{t+\tau,t} - \pi^*)^2/T, \quad (7)$$

$$\mathrm{MSG}_t^{u} \equiv \sum_{\tau=0}^{T} (u_{t+\tau,t} - u^*)^2/T. \quad (8)$$

Thus, the MSG for a variable is the average deviation of the forecast of the future variable from the target for the variable. A smaller MSG for a variable indicates better mandate fulfillment for the variable, with a zero MSG indicating (unlikely) perfect mandate fulfillment.[14]

Given this, the rule for the FOMC that would best fulfill its mandate is to select a policy rate path, i^t, so that that, conditional on this path and current information about the economy and prices, the shocks hitting the economy, and the transmission mechanism of monetary policy, the resulting forecasts for inflation, π^t, and unemployment, u^t, "look good." Here, "looking good" means "mandate

[14]Division by the horizon T to get mean squared gaps instead of cumulative squared gaps is not necessary but allows a convenient analogy with the well-known concept of mean squared errors in statistics.

consistent," in the sense of mitigating the deviations of the inflation and unemployment forecasts from, respectively, the inflation target and the long-run sustainable unemployment rate; more precisely, minimizing the sum of the MSGs of inflation and unemployment, (6). Equal weight on the MSGs indicates a "balanced approach."

Forecasts of inflation and unemployment can be generated with the methods of anticipated alternative policy rate paths of Laséen and Svensson (2011) or of unanticipated policy interventions of Leeper and Zha (2003), or a combination of the two methods. Svensson (2005) shows how to incorporate judgment in a systematic way. In particular, determinacy in a forward-looking setting requires a terminal condition on a given policy rate path, which is further discussed in section 3.5.

Note that setting monetary policy to minimize the deviations of the *forecasts* from their targets means that the forecasts of inflation and unemployment are effectively used as *intermediate* target variables for inflation and unemployment, the actual target variables. Using forecasts as intermediate target variables justifies the name *forecast targeting.*[15]

This decisionmaking process means that the monetary policy decision takes into account all relevant new and old information available to the FOMC, including information about the economic activity and prices, the inferred shocks hitting the economy, and how the inflation and unemployment forecasts depend on the policy rate path, that is, the transmission mechanism of monetary policy.

More precisely, the decisionmaking process means that new information is "filtered through the forecasts," and such filtering determines what information is relevant for the decision. New information that for a given policy rate path affects the forecasts of inflation and unemployment is relevant for the decision; new information that does not affect the forecasts for a given policy rate path is not relevant for the decision.

The policymakers need to have a view or a "model" of the transmission mechanism and the determination of future inflation and

[15]The idea that inflation targeting implies that the inflation forecast becomes an intermediate target was introduced in King (1994). The term "inflation-forecast targeting" was introduced in Svensson (1997), and the term "forecast targeting" in Svensson (2003b).

unemployment. But this does not have to be a singular model such as the Federal Reserve's FRB/US or the Riksbank's RAMSES. Instead it is a general view and understanding of the working of the macroeconomy and monetary policy, inspired by a set of several models and a large set of research results. It is the kind of general understanding of the macroeconomy that most senior and experienced macroeconomists would have. It is also the understanding that any particular issue may be well represented by a particular model, but every model is a simplification for a particular purpose, and no single model is enough for all purposes. In particular, forecasts and policy decisions cannot rely on models and simple observable data alone. All models are drastic simplifications of the economy, and data give an imperfect view of the state of the economy. Therefore, judgmental adjustments in both the use of models and the interpretation of their results—adjustments due to information, knowledge, and views outside the scope of any particular model—are a necessary and essential component in modern monetary policy (Svensson 2005).

Furthermore, this decisionmaking process involves continuing updating and learning about the state and working of the economy and the transmission process of monetary policy. Indeed, I would like to argue that this decisionmaking process is fully consistent with what is called Bayesian learning and Bayesian optimal policy. A Bayesian optimal policy involves in this context not only choosing a policy rate path (and any other policy instruments) so as to minimize an intertemporal loss function, conditional on all relevant prior and new information, including all information about the state of the economy and the outlook for relevant exogenous variables, as well as continuous learning by Bayesian signal extraction and updating. It also includes taking into account the different possible models of the transmission mechanism and the probabilities that they are correct, other aspects of model uncertainty, judgment, scientific evidence, practical experience, and so on. Indeed, such Bayesian optimal policy is arguably the most robust monetary policy among available alternatives.[16]

[16]See the discussion in Svensson (2013) of the robust interest rules promoted in Orphanides and Wieland (2013) for more about the properties of forecast targeting and Bayesian optimal policy.

3.2 An Example of Forecast Targeting versus the Taylor Rule

An example from Svensson (2005) can illustrate the importance of taking all relevant information into account. In this example, the new relevant information is information about future inflation and demand shocks. Two empirical models of the U.S. economy are used.

3.2.1 A Backward-Looking Model

The Rudebusch and Svensson (1999) backward-looking estimated model of the U.S. economy has two equations for inflation and the output gap. By assuming an Okun coefficient of 2, the output gap equation can be converted to an unemployment gap equation, using the simplified relation $u_t = -\,y_{t/2}$, where u_t and y_t here denote the unemployment gap and output gap, respectively, measured in percentage points. Then the equations for the inflation rate and the unemployment gap are (with converted estimates rounded to two decimal points)

$$\pi_{t+1} = 0.70\,\pi_t - 0.10\,\pi_{t-1} + 0.28\,\pi_{t-2} + 0.12\,\pi_{t-3} - 0.28\,u_t + z_{\pi,t+1} \tag{9}$$

$$u_{t+1} = 1.16\,u_t - 0.25\,u_{t-1} + 0.05\,(\Sigma_{j=0}^{3} i_{t-j}/4 - \Sigma_{j=0}^{3}\pi_{t-j}/4) + z_{u,t+1}. \tag{10}$$

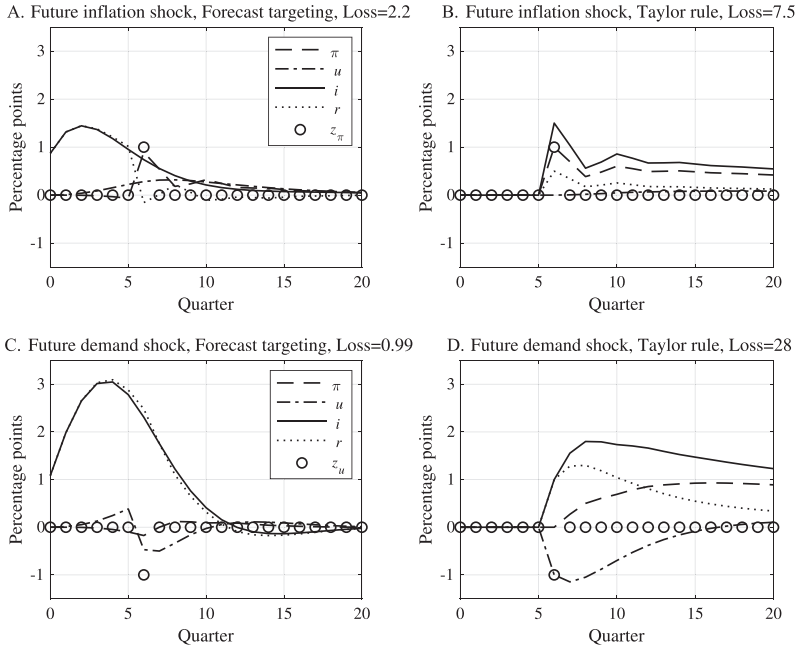
The period is a quarter, π_t denotes quarterly GDP inflation measured in percentage points at an annual rate, and the policy rate, i_t, is the quarterly average of the federal funds rate, measured in percentage points at an annual rate. All variables are measured as differences from their means, their steady-state levels. Furthermore, $z_{\pi,t+1}$ and $z_{u,t+1}$ denote shocks to inflation and the unemployment gap in quarter $t+1$.

The period loss function used in the optimization is

$$L_t = \pi_t^2 + u_t^2 + 0.2(i_t - i_{t-1})^2, \tag{11}$$

where π_t inflation is measured as the difference from the inflation target, so the steady state of inflation is assumed to equal the inflation target. The discount factor, δ, is set to unity. As in Rudebusch and Svensson (1999), a modest weight on interest rate smoothing

Figure 1. Forecast Targeting and the Taylor Rule: Backward-Looking Model



Notes: The period loss function is $L_t = \pi_t^2 + u_t^2 + 0.2(i_t - i_{t-1})^2$. The Taylor rule is $i_t = 1.5\,\pi_t - u_t$. The real interest rate that affects the unemployment gap in the Rudebusch and Svensson (1999) model is $r_t = i_t - \pi_t$, which is plotted in the figure. The intertemporal loss reported excludes the interest rate term, which is small. See Svensson (2005) for details.

is added as a convenient way not to get unrealistically large policy rate movements.[17] Standard impulse responses to current shocks are shown in figure A.1 in the appendix.

Figure 1 shows a situation in which, before and including quarter 0, inflation has equaled the target, the unemployment gap has equaled zero, and all inflation and demand shocks have been zero.

[17]Using the quarterly inflation rate at an annual rate is a simplification. Using a four-quarter inflation rate would be closer to the Federal Reserve's mandate. The main points of the example are not affected by the simplification.

Furthermore, the central bank's expectations—that is, its probability means—of future shocks have all been zero (although the perceived variance of the shocks may be significant).

Panel A refers to a situation in which the central bank receives new information in quarter 0 that makes it expect an inflation shock in quarter 6 with a mean of 1 percentage point, whereas expected inflation shocks for all other quarters and expected unemployment shocks for all quarters remain zero. The circles in panel A show the expected inflation shocks. Behind these expected shocks could be a perceived probability distribution of future realizations with substantial variance.

The central bank is assumed to conduct optimal forecast targeting, taking into account the new information about a future expected shock to inflation. Given this, panel A has two interpretations. In the first interpretation, it shows the central bank's expectations in quarter 0 of the future inflation shocks, $z_\pi^0 \equiv \mathrm{E}_0(z_{\pi 1}, z_{\pi 2}, \ldots, z_{\pi T}) \equiv (z_{\pi 1,0}, z_{\pi 2,0}, \ldots, z_{\pi T,0}) \equiv \{z_{\pi\tau,0}\}_{\tau=1}^T$, as well as the central bank's optimal policy rate path in quarter 0, $i^0 \equiv \{i_{\tau,0}\}_{\tau=0}^T$, and the corresponding optimal policy projection for inflation, $\pi^0 \equiv \{\pi_{\tau,0}\}_{\tau=0}^T$, unemployment, $u^0 \equiv \{u_{\tau,0}\}_{\tau=0}^T$, and real interest rate, $r^0 \equiv \{r_{\tau,0}\}_{\tau=0}^T$.[18] In the second interpretation, it shows the impulse responses starting in quarter 0 to an expected positive inflation impulse in quarter 6. That is, it shows the time series of the realizations of the policy rate, inflation, and unemployment for the particular realizations of the future inflation and unemployment shocks that are exactly equal to the central bank's expectations in quarter 0.

Using the second interpretation, we see that the central bank raises the policy rate preemptively from quarter 0, in order to contain the future inflation shock. This causes the unemployment gap to increase somewhat, but inflation is contained. When the inflation shock arrives in quarter 6, inflation jumps 1 percentage point but falls back toward the target in a couple of quarters. The unemployment gap slowly falls back to zero. The total intertemporal loss, (5), is 2.2 units (including only the inflation and unemployment terms

[18] In the Rudebusch and Svensson (1999) model, the real interest rate that affects the unemployment gap and is plotted in figure 1 is $r_t = i_t - \pi_t$, that is, the policy rate minus the current inflation rate.

and excluding the term corresponding to interest rate smoothing, which is in any case small in these simulations).

Panel B refers to a situation in which the central bank just follows the standard Taylor rule, (1), suggested by the CHOICE and FORM Acts.[19] Expressed in terms of the unemployment gap, under the maintained assumption of an Okun coefficient of 2, it is

$$i_t = 1.5\,\pi_t - u_t. \tag{12}$$

Under the Taylor rule, the central bank does not move the policy rate until inflation or unemployment moves. It raises the policy rate by 1.5 percentage points in quarter 6 when the inflation shock hits. But the Taylor rule puts the central bank behind the curve. It only slowly increases unemployment and slowly brings inflation back toward the target. The loss is 7.5 units, substantially larger than under forecast targeting.[20] This is due to both the lack of preemptive policy with the Taylor rule and the modest tightening that the Taylor rule implies once the inflation shock occurs. The policy rate is only raised a modest amount above the increased inflation, and the real interest rate increases only a little. In contrast, in panel A, the policy rate is raised substantially before the shock while inflation is flat, so the real interest rate rises substantially. The backward-looking Rudebusch-Svensson model has substantial inertia and autocorrelation in both inflation and unemployment, and optimal policy involves strong preemptive movements to keep inflation close to target and the unemployment gap close to zero.

Panel C shows the situation in which the central bank in quarter 0 instead receives new information about a future positive demand shock in quarter 6, which is assumed to result in an expected negative unemployment shock of 1 percentage point. Optimal preemptive policy involves a substantial tightening and rise of the real interest rate that succeeds in stabilizing inflation close to the target as well as stabilizing the unemployment gap movements. The loss is 1.0 unit.

[19] Using the quarterly inflation rate at an annual rate instead of the four-quarter inflation rate in the Taylor rule is a simplification but does not affect the main points of the example.

[20] The losses reported include any losses beyond the quarter-20 forecast horizon.

Panel D shows that, in this case, the Taylor rule raises the real interest later and less, which stabilizes inflation and unemployment much less than forecast targeting. The lack of preemptive policy before the shock and the modest effectiveness once the shock has occurred allows inflation to rise and stay above the target for a considerable time. Again, the central bank gets behind the curve. The loss is a large 28 units.

What if the expected shock does not materialize? Consider the top two panels, showing the impulse responses to the new information in quarter 0 about the expected future inflation shock in quarter 6. Suppose there are no realizations of current shocks and no further information about expected future shocks through quarter 5. Then the economy would develop according to the panels through quarter 5. Suppose then, in (the beginning) of quarter 6, that the expected shock does not materialize.

The shock not materializing is equivalent to receiving the expected inflation shock of 1 percentage point but also at the same time an unanticipated inflation shock of -1 percentage point. Under forecast targeting in panel A, the appropriate policy response is equal to the negative of the policy rate impulse response of a current 1 percentage point inflation shock. It is the negative of the impulse response shown in panel A in figure A.1. Thus, the negative of the impulse responses in figure A.1 is shifted seven quarters to the right, to quarter 6, and then added to the impulse responses to the expected future shock in panel A of figure 1.

For the Taylor rule in panel B of figure 1, the unexpected negative inflation shock cancels the expected shock, and neither inflation, GDP, nor the policy rate changes. Put differently, the impulse responses in panel B starting in quarter 6 are equal to the negative policy responses starting in quarter 0 in panel B of figure A.1.

Similarly, in panel C of figure 1, if the expected demand shock does not materialize, the response under forecast targeting are given by adding the appropriately scaled and shifted negative of the impulse responses in panel C of figure A.1.

This is a very special case of an expected future shock and later a current unexpected shock that exactly cancels the expected shock. More realistically, there would be new information about current and expected future shocks every quarter. For instance, expectations

about future shocks in a particular future quarter would be updated over time. Forecast targeting would adjust the policy rate path and the forecasts of inflation and GDP, taking into account both current and expected future shocks. The Taylor rule would just respond to shocks that affect current inflation and GDP.

3.2.2 A Forward-Looking Model

The forward-looking estimated New Keynesian model of Lindé (2005) of the U.S. economy also has two equations for the inflation rate and the output gap. With the assumption of an Okun coefficient of 2, it can be expressed in terms of the unemployment gap, with the following (converted) parameter estimates,

$$\pi_t = 0.457\,\pi_{t+1|t} + (1 - 0.457)\,\pi_{t-1} - 0.096\,u_t + z_{\pi t}, \tag{13}$$

$$u_t = 0.425\,u_{t+1|t} + (1 - 0.425)\,u_{t-1} + 0.078\,(i_t - \pi_{t+1|t}) + z_{ut}, \tag{14}$$

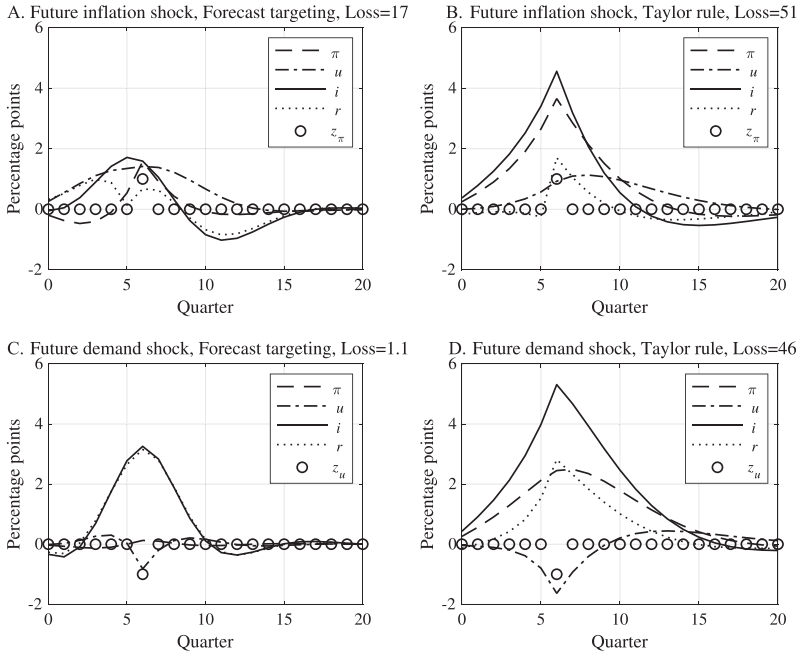
where $\pi_{t+1|t}$ and $u_{t+1|t}$ denote private-sector rational expectations in quarter t of inflation and unemployment in quarter $t+1$. Standard impulse responses to current shocks are shown in figure A.2 in the appendix.

Panel A in figure 2 shows the same situation for the central bank as in panel A in figure 1. The central bank has in quarter 0 received new information that results in an expected positive inflation shock of 1 percentage point in quarter 6. Furthermore, the private sector has received the same information about the future shock, either independently or conveyed by the central bank, and has the same expectations as the central bank.

Optimal forecast targeting again involves preemptive policy. Policy tightening raises the unemployment gap and shortens the period of inflation overshooting the target. The loss is 17 units.

Panel B shows the outcome under the Taylor rule. Because the private sector is forward looking and takes the future inflation shock into account, inflation increases substantially above the target before the shock arrives. Thus, the Taylor rule raises the policy rate and involves some indirect preemptive policy. However, the private sector is even more preemptive in raising inflation substantially, and the real interest rate does not move until the inflation shock occurs,

Figure 2. Forecast Targeting and the Taylor Rule: Forward-Looking Model



Notes: Forecast targeting minimizes the loss function $L_t = \pi_t^2 + u_t^2 + 0.2(i_t - i_{t-1})^2$. The Taylor rule is $i_t = 1.5\,\pi_t - u_t$. The real interest rate that affects the unemployment gap in the Lindé (2005) model is $r_t = i_t - \pi_{t+1|t}$, which is plotted in the figure. The intertemporal loss reported excludes the interest rate term, which is small. See Svensson (2005) for details.

after which inflation falls.[21] The central bank is again behind the curve, the policy tightening in terms of the real interest rate is too late and too small, and a large overshoot of inflation results. The loss is a large 51 units.

Panel C again shows the situation in which the central bank and the private sector have received information about a positive demand shock and a corresponding expected negative unemployment shock of 1 percentage point in quarter 6. The optimal policy projection,

[21] Recall that the forecast in quarter 0 of the real interest rate in quarter t is $r_{t,0} = i_{t,0} - \pi_{t+1,0}$, the forecast of the policy rate in quarter t minus the forecast of the inflation rate in quarter $t+1$.

taking this information into account, is to raise the policy rate before the expected unemployment shock, which moderates the expected impact on the unemployment gap. The inflation projection remains very flat, and the real interest rate is close to the policy rate. The resulting intertemporal loss is very small, 1.1.

Panel D shows that the private sector reduces the unemployment gap and increases inflation in anticipation of the future demand shock. The Taylor rule then raises the policy rate substantially. The increase in the real interest rate is also substantial. But the increase is nevertheless too late compared to forecast targeting and as a result the central bank is also here behind the curve. A large overshoot of inflation and a substantial undershoot of the unemployment gap result. The loss is large, 46 units, especially in comparison with the small loss from the unemployment shock under forecast targeting.

As discussed above for the backward-looking model, the case when the expected shock does not materialize is equivalent to an unexpected negative shock that cancels the expected shock. The outcome can be shown by adding the appropriately shifted negative impulse responses in figure A.2 to figure 2. The loss is less if the expected shock does not materialize. The Taylor rule does much worse than forecast targeting, regardless of whether the expected shock materializes or not.

In summary, these results indicate that the Taylor rule performs considerably worse than forecast targeting when there are expected future shocks.

In these two examples, the optimal policy rate path and the forecasts of inflation and unemployment under forecast targeting are for simplicity generated by two given models, where information about expected future shocks are taken into account. In practical applications of forecast targeting, the policy rate path and forecasts of the target variables would normally be generated by a combination of model simulations from several models as well as from a substantial application of judgment—that is, information, knowledge, and views outside the scope of a particular model.

3.3 Implementation

The decision under forecast targeting selects the policy rate path and inflation and unemployment forecast that, *if believed by the market*

and other economic agents, best fulfills the mandate. A successful implementation of the selected policy involves making the policy rate path credible, in the sense of market expectations aligning with the policy rate path. The policy rate path can be seen as representing the FOMC's *intended* monetary policy, or *intended* financial conditions. Market expectations of future policy rates and resulting market yield curves for different assets can be seen as the *actual* monetary policy, the *actual* financial conditions. The implementation of monetary policy involves trying to make the actual financial conditions equal to the intended financial conditions.

However, not only market expectations of the future policy rate but also economic agents' expectations of future inflation, unemployment, GDP, and other economic variables matter. In particular, if the FOMC manages to make the inflation target credible, in the sense of making economic agents' inflation expectations align with the inflation target, stabilization of inflation around its target is much easier, because economic agents' individual decisions that result in (economy-wide) inflation are much influenced by the agents' expectations of (economy-wide) inflation.[22]

The most effective contribution to making the policy rate path and inflation forecasts credible is to publish them and justify the decision. In addition to justifying how new information since the last decision has affected the forecasts and the selected policy rate path, the justification of the decision may include the publication of inflation and unemployment forecasts for alternative policy rate paths different from the selected one and the demonstration that these forecasts do not fulfill the mandate to the same extent. That demonstration may use MSGs for inflation and unemployment as quantitative measures of the degree of mandate fulfillment.

3.4 The Forecast-Targeting Rule Summarized

The forecast-targeting rule can be summarized as these three steps:

[22]See Svensson and Woodford (2005) for details.

(i) For a given initial policy rate path (for example, the policy rate path from the previous decision), construct new inflation and unemployment forecasts, taking into account new information received since the previous decision.

(ii) If the new inflation and unemployment forecasts "look good" (meaning they best fulfill the mandate), select the given policy rate path as the decision; if the new inflation and unemployment forecasts do not look good, adjust the policy rate path so that they do look good.

(iii) Publish the policy rate path and inflation and unemployment forecasts and justify the decision in order to make the published path and forecasts credible, meaning making market participants' and other economic agents' expectations align with the published path and forecasts. The justification of the decision may include the publication of inflation and unemployment forecasts for alternative policy rate paths different from the selected one and the demonstration that these forecasts do not fulfill the mandate to the same degree. MSGs for inflation and unemployment as quantitative measures of the degree of mandate fulfillment may be used.

3.5 Determinacy

By the well-known result of Sargent and Wallace (1975), for an exogenous policy rate the (rational-expectations) equilibrium in a forward-looking model may be indeterminate. It is therefore necessary to make sure that any selection and announcement of a policy rate and policy rate path under forecast targeting with the help of a forward-looking model do not encounter the problem of indeterminacy.

Uniqueness of policy simulations with exogenous policy rate paths in forward-looking models can be ensured by two alternative terminal conditions at a future terminal horizon at quarter $\bar{T}$, beyond the horizon at quarter T for the published forecasts. One terminal condition is that policy switches to a given reaction function for which the equilibrium is unique—for example, the optimal reaction function without any judgmental modifications.

Alternatively, the forecasts of inflation and unemployment are restricted to reach a steady state in which they are equal to, respectively, the inflation target and the long-run sustainable unemployment rate, that is, $\pi_{t+\bar{T},t} = \pi^*$ and $u_{t+\bar{T},t} = u^*$. The terminal horizon is then extended until the forecasts for inflation and unemployment within the forecast horizon have converged and are no longer sensitive to the terminal horizon (see Svensson 2005; Svensson and Tetlow 2005; Laséen and Svensson 2011).

However, there is a further indeterminacy issue, which is discussed in detail in Svensson and Woodford (2005) and summarized in Svensson (2011, section 3.7). The announcement of a policy rate and a policy rate path (also if they are consistent with an optimal reaction function) and corresponding forecasts of inflation and unemployment may, even if these are credible with economic agents and their expectations are aligned with the announcement, not be sufficient to ensure determinacy. This can be understood by noting that the optimal policy rate under commitment will be a function of present and past exogenous shocks and thereby in the end be exogenous and subject to the Sargent and Wallace (1975) result. In this case, an explicit or implicit out-of-equilibrium commitment—understood by and credible with the economic agents—may be required. Such a commitment is typically quite intuitive, such that it is understood that the central bank will—all else being equal—raise (lower) the actual policy rate sufficiently above (below) the previously announced policy rate path, if realized inflation exceeds (falls short of) the forecast or realized unemployment falls short of (exceeds) the long-run sustainable unemployment rate. Such a commitment can thus be seen as a kind of Taylor principle applied to realized deviations of inflation and unemployment from the previous forecasts.[23]

3.6 Time Consistency

There is another somewhat technical issue that needs to be sorted out, namely time consistency. In forward-looking models, or more generally in situations in which economic agents' decisions depend

[23]Similar results as in Svensson and Woodford (2005) have later appeared in Atkeson, Chari, and Kehoe (2010).

on their expectations of future outcomes, there is a well-known time-consistency problem. The time-consistency problem implies that optimization under commitment, that is, optimization under commitment to a future history-dependent policy, is normally better than optimization under discretion, that is, reoptimization in the future without any such commitment (Backus and Driffill 1986; Currie and Levine 1993). Woodford (1999) has suggested optimization "in a timeless perspective" as a possible solution to this.

The issue of time consistency in this context is discussed in detail and resolved in Svensson and Woodford (2005) and summarized in Svensson (2011, section 3). The desired history dependence under commitment can be imposed in two ways. First, the intertemporal forecast loss, (5), can be modified by the addition of a term that represents the cost of deviating from previously announced policy. The MSGs, (7) and (8), can then be adjusted by adding to each MSG this term divided by $2T$.[24] Alternatively, as shown in Giannoni and Woodford (2003) and Svensson and Woodford (2005) and summarized in Svensson (2011), a history-dependent restriction on the policy rate path and the forecasts can be added. This means that (5) is minimized for a restricted set of policy rate paths and forecasts that satisfy this restriction in addition to the equations of the model used; see Svensson (2011, equations (28) and (29)). If the FOMC decides to restrict its policy choices to those consistent with

[24]The intertemporal forecast loss, (5), is then replaced by

$$\mathcal{L}_t = \sum_{\tau=0}^{T} (\pi_{t+\tau,t} - \pi^*)^2 + \sum_{\tau=0}^{T} (u_{t+\tau,t} - u^*)^2 + \ell_t, \tag{15}$$

where ℓ_t is the cost of deviating from previous promises, more precisely a history-dependent function of the difference between the quarter-t realization of the forward-looking variables and the previous forecasts and expectations of these variables; see Svensson (2011, equation (26)). Then the definition of the MSGs, (7) and (8), is replaced by

$$\mathrm{MSG}_t^{\pi} \equiv \sum_{\tau=0}^{T} (\pi_{t+\tau,t} - \pi^*)^2 / T + \ell_t/(2T), \tag{16}$$

$$\mathrm{MSG}_t^{u} \equiv \sum_{\tau=0}^{T} (u_{t+\tau,t} - u^*)^2 / T + \ell_t/(2T). \tag{17}$$

such commitment, the Federal Reserve staff would then present policy alternatives that either have modified MSGs or are subject to the restriction mentioned.

Alternatively, the FOMC may decide that commitment and related history dependence is unenforceable and impractical. Then the policy simulations can be done under the assumption of discretion, as discussed in Svensson (2011, section 3.6). It may also be relevant to show policy simulation under both commitment and discretion, so as to examine how different they are and whether the difference is of any practical significance.

The transparency of forecast targeting, with the publication, explanation, and justification of policy rate paths and forecasts of inflation and unemployment, should allow the FOMC a substantial degree of commitment, if it becomes established that deviations from previously published paths and forecasts only come with good explanations. Forecast targeting may imply a policy that is approximately optimal under commitment.

In a situation in which the policy rate is restricted by the ELB, a commitment to a policy rate that is "lower for longer" is normally more effective in stimulating the economy by lowering longer-term interest rates. Such a commitment may involve a significant time-consistency problem. Enforcing a commitment in such situations is discussed in some detail in Bernanke (2017), including the role of a temporary price-level target path.

3.7 The Reaction Function

It is common to argue that central banks should convey their reaction function to the market participants and other economic agents. However, under forecast targeting the reaction function, meaning how the policy rate and the policy rate path respond to information available to the central bank, is far too complex to write down as a simple formula such as a Taylor-type rule. It is actually too complex to write down, period—that is, as a mathematical function describing precisely how the policy rate responds to different kinds of information.

First, the reaction function is not just the current policy rate (a scalar) that is a function of a list of arguments. It is the current policy rate *and* the policy rate path (a vector) that is a function of

a list of arguments. Second, an explicit list of arguments, consisting of the possible pieces of information that the policy rate and policy rate path may need to respond to, is too long to be conveyed. In particular, central banks cannot anticipate all future pieces of information and shocks that may occur and may be relevant to respond to. In terms of a model, the optimal reaction function will require responses to all the relevant state variables and the expected future shocks. In the real economy this is a very long and changing list. Third, it is impossible to specify in advance the appropriate magnitude of all the response coefficients, in particular for any new shocks appearing. Furthermore, the response coefficients will depend on specific circumstances at the time, including how persistent the shocks are judged to be, whether the transmission mechanism of monetary policy is judged to be weaker or stronger than usual, and so on.

As an example, even for the relatively simple completely specified backward-looking model with just two equations and information that only affects expected future shocks to inflation and unemployment, (9)–(11), the optimal reaction function, taking into account expected future shocks, is rather complicated. It is given by

$$i_t = F_i X_t + R_i z^t, \tag{18}$$

where F_i is the row-9 vector of response coefficients to the column-9 vector of predetermined variables, $X_t \equiv (\pi_t, \pi_{t-1}, \pi_{t-2}, \pi_{t-2}, u_t, u_{t-1}, i_{t-1}, i_{t-2}, i_{t-3})'$.[25] Furthermore, $R_i z^t$ represents the response to the expected future shocks, $z^t = \{z_{t+1+\tau,t}\}_{\tau=0}^{\infty} \equiv \{\mathrm{E}_t z_{t+1+\tau}\}_{\tau=0}^{\infty}$, where $z_t = (z_{\pi t}, z_{ut})'$. It is given by the first element of the column-10 vector resulting from the linear operator R given by

$$Rz^t = \sum_{\tau=0}^{\infty} R_\tau \begin{bmatrix} \mathrm{E}_t z_{t+1+\tau} \\ 0_{7\times 1} \end{bmatrix} = \sum_{\tau=0}^{\infty} R_\tau \begin{bmatrix} z_{t+1+\tau,t} \\ 0_{7\times 1} \end{bmatrix}, \tag{19}$$

where the 10×9 matrices $\{R_\tau\}$ satisfy $R_\tau = HJ^\tau K$, $\tau = 0, 1, ...,$ where the matrices H, J, and K are outputs of the algorithm that solves for the optimal policy and depend on the parameters of the

[25] In this case, $F_i = (1.215, 0.408, 0.516, 0.167, -2.772, 0.673, 0.474, -0.064, -0.033)$.

model and the loss function (see Svensson 2005, online appendix, for details).[26]

However, fortunately the reaction function can be conveyed in more general and approximate but still both systematic and simple terms. For example, if initially the forecasts look good, for any new piece of information that shifts the inflation forecast up (down) and/or shifts the unemployment forecast down (up) for a given policy rate path, policy will normally be tightened (eased), meaning that the policy rate path will shift up (down). If this response is understood by and credible with the market participants, any new information that is deemed to shift up (down) the inflation outlook or shift down (up) the unemployment outlook may result in a market response that shifts up (down) the yield curve (or, more precisely, a forward-rate curve such as the overnight index swap (OIS) curve), even before the central bank has responded with the same shift in the policy rate path.

Thus, with such behavior of the market, the market may do a good part of the work of the central bank before the central bank acts itself. In that case, it is of course important for maintaining credibility that the central bank completes the policy move by moving its policy rate and policy rate path accordingly at the next decision. The publication and justification of the new policy rate path and the forecasts will also be an opportunity to try to correct any over- or under-adjustment by the market. For this, it may be effective to let the justification include how the central bank interprets how new information has shifted the forecasts of inflation and unemployment before any adjustment of the policy rate path (corresponding to step 1 in the summary of the forecast-targeting policy rule in section 3.4).[27]

In this case, to a naive observer, it may look like the central bank is to a large extent just following the market, whereas the truth is

[26]For expected inflation shocks up to six quarters ahead, the response coefficients are $\{\partial i_t/\partial z_{\pi,t+1+\tau,t}\}_{\tau=0}^{5} = (1.144, 1.083, 1.031, 0.983, 0.933, 0.872)$. For the expected unemployment shocks, they are $\{\partial i_t/\partial z_{u,t+1+\tau,t}\}_{\tau=0}^{5} = (-2.671, -2.514, -2.268, -1.926, -1.508, -1.083)$. Thus, $i_0 = 0.872$ in panel A of figure 1 and $i_0 = 1.083$ in panel C (in the latter case the expected shock is negative in quarter 6).

[27]Federal Reserve Board (2019, p. 39) makes a very similar point.

that a well-informed market to a large extent is anticipating the central bank's policy setting.

The central bank can convey the reaction function in these more general terms more explicitly by showing model simulations for alternative assumptions about the shocks and exogenous variables. For example, by showing a simulation with a larger future shock to inflation, the central bank can demonstrate that the optimal policy rate path shifts up (as in the example in section 3.2). Similarly, by showing a simulation with a larger negative future shock to unemployment, it can show that the optimal policy rate path shifts up also for this shock.

Such policy simulations showing optimal policy rate paths and outcomes for alternative assumptions about exogenous variables and shocks can be called "alternative scenario" simulations. Their purpose is thus to illustrate how the central bank would respond to alternative exogenous shocks and new information.

It is worth noting that alternative-scenario simulations are conceptually different from simulations showing the outcome of (forecasts for) inflation and unemployment for alternative policy rate paths for given assumptions about the exogenous variables and shocks. These can be called "alternative policy" simulations. Their purpose is to justify the selected optimal policy rate path by showing that alternative policy rate paths—for example, tighter or looser policy—would be expected to lead to less good outcomes than the one selected. MSGs can be used for this purpose (section 4.1).

3.8 Forecast Targeting, Instrument Rules, and Targeting Rules

A monetary policy rule can generally be defined as "a prescribed guide for monetary policy conduct."[28] Forecast targeting, summarized in a simplified way in section 3.4, is thus a monetary policy rule. However, most of the literature on monetary policy rules have used a more narrow interpretation of a policy rule, namely what can be called an *instrument rule*, in which the central bank's policy instrument, typically a short interest rate, is set as a given function of a

[28]Svensson (2003b) provides a more extensive discussion of monetary policy rules.

given set of observable variables. In a given model, the corresponding optimal reaction function—for example, (18) and (19)—can be seen as an optimal instrument rule. But the discussion has mostly focused on *simple* instrument rules, in which the policy instrument is set as a function of only a few variables. The best known is the Taylor (1993) rule, shown in (1). Similar simple instrument rules can be called Taylor-type rules.[29]

However, other rules are possible, such as *targeting rules* (also known as target rules), that is, conditions for (the forecasts of) the target variables. One such rule is that policy should be set such that the inflation forecast is close to the inflation target at some specified horizon, such as two years. This rule is arguably too rigid to represent flexible inflation targeting and the dual mandate (Svensson 1997). Another rule—consistent with the dual mandate—is that the

[29]Instrument rules can be divided into two categories, explicit instrument rules and implicit instrument rules (Svensson 2003b; Svensson and Woodford 2005). An explicit instrument rule is a reaction function where the instrument responds to predetermined variables only. Its implementation then consists of the central bank observing the predetermined variables in the beginning of the period, and then calculating, announcing, and setting the instrument according to this instrument rule. The implementation obviously requires that the relevant predetermined variables must be observed by the central bank, but since the predetermined variables in a particular period are independent of the instrument setting in that period, no further complications arise. An implicit instrument rule is a relation between the current instrument and some of the current forward-looking variables. Then, since the forward-looking variables depend on the instrument setting, the instrument and the forward-looking variables are simultaneously determined. Thus, an implicit instrument rule is actually an equilibrium condition, a relation that holds in equilibrium. The implementation of an implicit instrument rule, that is, how to get to the desired equilibrium, is not trivial but a complex issue. This fact has largely been overlooked in the literature, except, for example, in Svensson (2003b, 2005) and Svensson and Woodford (2005).

In the forward-looking model in section 3.2.2, the Taylor rule in (12) implies that the policy rate responds to the forward-looking variables inflation and unemployment, which in turn respond to the policy rate. That is, the Taylor rule is an implicit instrument rule and really a sophisticated equilibrium condition. A more realistic Taylor rule is when the policy rate responds to the lagged inflation and unemployment gap, making it an explicit instrument rule in the forward-looking model,

$$i_t = 1.5\,\pi_{t-1} - u_{t-1}. \tag{20}$$

In the forward-looking model above, the explicit Taylor rule performs slightly worse than the implicit one.

forecasts of the inflation gap and the unemployment gap should normally have the same sign and be in reasonable proportion to each other until they close (Qvigstad 2005).

An optimal targeting rule is a first-order condition for optimal monetary policy that involves the targeting variables only.[30] It corresponds to the standard efficiency condition of equality between the marginal rates of substitution and the marginal rates of transformation between the target variables, the former given by the monetary policy loss function, the latter given by the transmission mechanism of monetary policy. An optimal targeting rule is invariant to everything else in the model, including additive shocks and judgment and the stochastic properties of additive shocks. Thus, it is a more compact and robust representation of optimal monetary policy, more robust than the optimal reaction function.

Targeting rules are called "target criteria" in Giannoni and Woodford (2003, 2017). They show that, with a quadratic loss function, it is always possible to rewrite the first-order condition for optimal policy as an optimal targeting rule that only contains target variables. As a well-known example, the standard New Keynesian model with only inflation and the output gap in the quadratic loss function has the simple and elegant targeting rule

$$\pi_t - \pi^* + \frac{\lambda}{\kappa}(y_t - y_{t-1}) = 0 \tag{21}$$

for optimal policy under commitment, where λ is the relative weight on stabilizing the output gap in the loss function and κ is the coefficient of the output gap in the New Keynesian Phillips curve.

The optimal targeting rule is simple and elegant in simple models, such as the forward-looking standard New Keynesian model or the simple backward-looking model of Svensson (1997).[31] They may therefore seem attractive as possible rules for monetary policy. However, in more realistic models with more lags, the optimal targeting rules become quite complicated linear combinations of several lags of (forecasts of) the target variables, with coefficients that depend

[30] Svensson (2003b), Svensson and Woodford (2005), and Giannoni and Woodford (2017) provide an extensive discussion of targeting rules.

[31] However, as noted in Svensson (2003b, section 5.4), they are quite different for simple backward- and forward-looking models.

in a more complicated way on the structure and coefficients of the models. In particular, deriving optimal targeting rules for the larger models actually used by central banks—such as the Federal Reserve's FRB/US or the Riksbank's RAMSES—would result in quite complicated targeting rules with many lags. Therefore, I consider optimal targeting rules from more realistic models non-operational, and I believe that it is so far unrealistic to use optimal targeting rules as guides for practical monetary policy, and that the simple and robust rule of forecast targeting is much preferable.[32]

4. Accountability

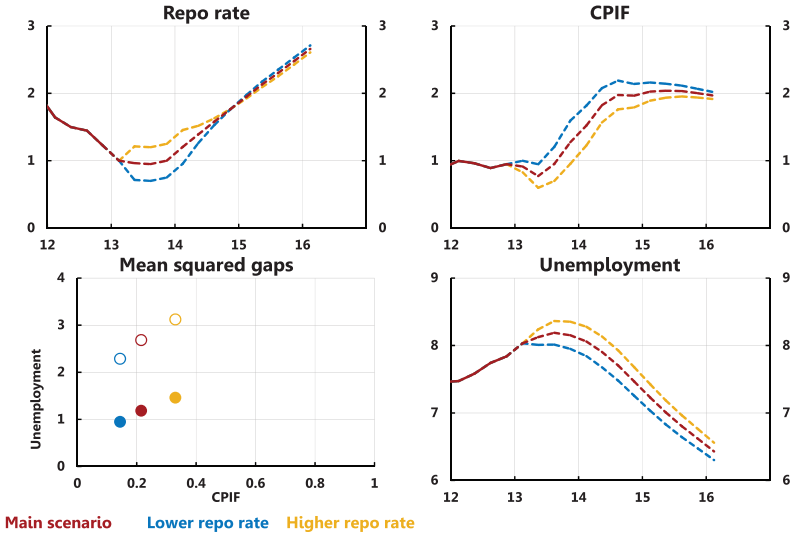
Can the FOMC be held accountable if it practices forecast targeting? Yes, forecast targeting can be scrutinized and reviewed by external observers if the Fed provides enough information. The Fed needs to publish the policy rate path and forecasts of inflation and unemployment and justify that they are internally consistent as well as consistent with available information about the economy and its structure and dynamics. It also needs to demonstrate that alternative policy rate paths—typically representing tighter and easier policy, respectively, than the selected policy—result in worse mandate fulfillment than the selected policy. These explanations, justifications, and demonstrations can be scrutinized and reviewed both in real time and after the fact—that is, after the outcome for inflation and unemployment have been observed—by external observers and experts and at the regular hearings in the Congressional oversight committees (Svensson 2012).

4.1 An Example: Reviewing the Policy Decision

The publication of the policy rate path and forecasts of inflation and unemployment allows a review of the policy decision, especially if the result from alternative policy rate paths is also published. An obvious criterion for an appropriate policy rate path is that it should

[32] With arguably not the best terminology, forecast targeting was called a "general targeting rule" and targeting rules were called "specific targeting rules" in Svensson (2003b).

Figure 3. Monetary Policy Alternatives around the Main Scenario



Source: Sveriges Riksbank (2013, figure 4).
Notes: For interpretation of the references to color in the text, the reader is referred to the web version of this article, available at http://www.ijcb.org. Empty circles indicate mean squared gaps calculated with long-run sustainable employment of 5.5 percent.

not be the case that a lower or higher policy rate path leads to better mandate fulfillment.[33]

It is possible to review this with the aid of a figure such as figure 3, which is from the minutes of the Riksbank policy meeting of February 2013 (Sveriges Riksbank 2013, figure 4). It is one of the similar four-panel figures that—during my term as a deputy governor and member of the Executive Board—I regularly brought to the Riksbank policy meetings for discussion and justification of my decision. The four panels show the repo rate path (top left; the repo rate is the Riksbank's policy rate), the MSGs for inflation and unemployment (bottom left), the forecast of the CPIF inflation rate

[33]This can be seen as a simple application of the so-called calculus of variations in optimization theory: A different policy should not give a better outcome.

(top right),[34] and the forecast of the unemployment rate (bottom right).

The center dashed red (dark gray in black-and-white print) lines refer to the majority's choice of a repo rate path and resulting forecasts of inflation and unemployment. The red (dark gray) circles in the bottom-left panel show the corresponding MSG points with the coordinates of the MSGs for inflation and unemployment measured along, respectively, the horizontal and vertical axes. The filled and unfilled circles refer to MSG points calculated with a long-run sustainable unemployment rate of, respectively, 6.25 percent (the majority's estimate) and 5.5 percent (my estimate). The inflation target is 2 percent.

The blue and yellow (medium and light gray) dashed lines refer to, respectively, a lower and a higher repo rate path and corresponding forecasts of inflation and unemployment. It is obvious that the lower repo rate path is better than the majority choice; the forecast of inflation is higher and closer to the target of 2 percent and the forecast of unemployment is lower and closer to the long-run sustainable unemployment rate, regardless of whether the latter is 6.25 percent or 5.5 percent. Consistent with this, in the bottom-left panel, the MSG points for the lower repo rate path (the blue [medium grey] circles) are southwest of the corresponding MSG points for the majority's repo rate path (the red [dark gray] circles).

In this trivial (but nevertheless real-world) example, the MSG points are not needed for the conclusion that the lower repo rate path is better than the center repo rate path. In this case, the lower policy rate path fulfills the mandate better for both inflation and unemployment; there is no tradeoff. It is also obvious that an even lower policy rate path would be better than the blue (medium gray) path in the top-left panel.

In a nontrivial case, there would be a tradeoff between stabilizing inflation and unemployment. For example, the forecasts of inflation and unemployment would both be above (or both be below) the inflation target and the long-run sustainable unemployment rate, respectively. Then the alternative MSG points for a given long-run sustainable unemployment rate would not line up southwest-northeast but

[34] CPIF inflation is CPI inflation when mortgage rates in the housing component of the CPI are held constant.

northwest-southeast, and the best repo rate path would be the one minimizing the sum of the MSGs.[35] Indeed, a necessary condition for a policy rate path to be a candidate for best fulfilling the mandate is that alternative policy rate paths do not result in MSG points southwest of the candidate's MSG point.[36,37]

Clearly, the publication of the policy rate path and corresponding forecasts of inflation and unemployment gives external observers and experts considerable possibilities to review how well the FOMC fulfills its mandate and to hold the FOMC accountable for this.

The discussion in Yellen (2012) of revolution and evolution in communication by central banks is very relevant here, in particular the figure of three policy rate paths and the resulting forecasts of inflation and unemployment shown in figure 4. In particular, a fourth panel could be added with the MSG points for the different policy rate paths.

5. Does the Federal Reserve Already Practice Forecast Targeting?

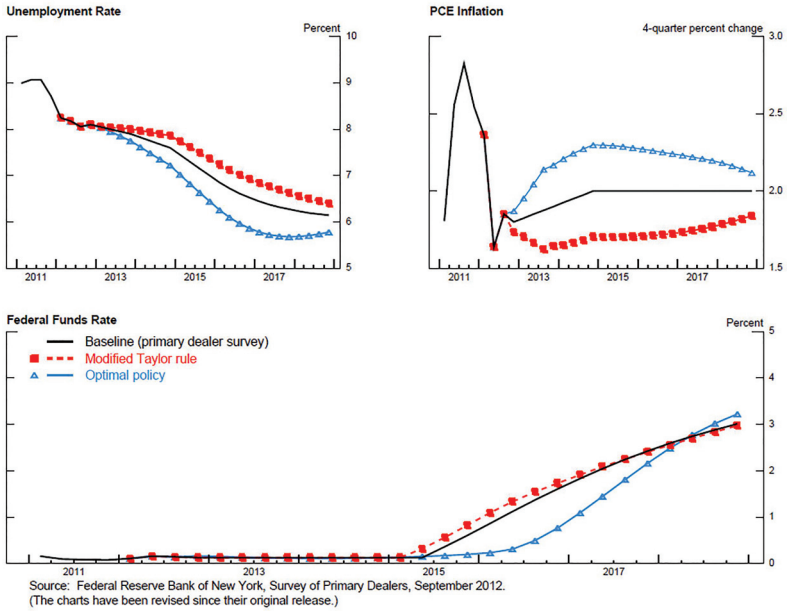
Forecast targeting can be summarized by the three steps in section 3.4. To what extent is the Federal Reserve already practicing

[35]ISO-loss lines with a slope of -1, corresponding to equal weight on the MSGs for inflation and unemployment, can be added to the bottom-left panel. Then the best repo rate path is the one for which the MSG point lies on the ISO-loss line that is closest to the origin.

[36]These MSGs were not adjusted for a possible time-consistency problem, discussed in section 3.6. There is no reason to believe that the adjustment would be large in this case or that it would be sufficiently different for the different policy alternatives to affect the ranking.

[37]Furthermore, in this real-world case, the majority's forecasts of inflation and unemployment in figure 3 were conditional on the majority's assumption of a high forecast of foreign interest rates, much above implied market forward rates. As shown in Sveriges Riksbank (2013, figure 5), assuming a lower forecast of foreign interest rates, in line with implied forward rates, resulted for a given policy rate path in a stronger exchange rate forecast and thereby an even lower inflation forecast and an even higher unemployment forecast. Then an even lower repo rate path was called for, which I dissented in favor of. See the minutes (Sveriges Riksbank 2013) for details. The minutes, published about two weeks after the meeting, are attributed; thus, in Sweden individual members of the Executive Board can be held accountable in real time not only for their votes and decisions but also for their individual statements and arguments at the policy meeting, regardless of whether they are dissenters or not.

Figure 4. Three Policy Paths: An Illustrative Exercise



Source: Yellen (2012).

forecast targeting? The Federal Reserve staff's optimal-control simulations described and discussed in Brayton, Laubach, and Reifschneider (2014) and used, for example, in Yellen (2012) and the optimal policy projections discussed in Svensson and Tetlow (2005) lend themselves well to steps (i) and (ii), the selection of an appropriate policy rate path. Regarding step (iii), the publication and justification of the decision, the FOMC is already publishing its Summary of Economic Projections (SEP), which includes economic projections of the FOMC participants under their individual assessments of projected appropriate monetary policy.

These projections have received considerable emphasis in the Chair's press conference after policy meetings. For example, as Chair Bernanke noted in his opening remarks at the press conference on April 27, 2011 (before the publication of interest rate projections, which began in January 2012):

> The Committee's economic projections provide important context for understanding today's policy action as well as the Committee's general policy strategy. Monetary policy affects output and inflation with a lag, so current policy actions must be taken with an eye to the likely future course of the economy. Thus the Committee's projections of the economy, not just current conditions alone, must guide its policy decisions. The lags with which monetary policy affects the economy also imply that the Committee must focus on meeting its mandated objectives over the medium term, which can be as short as a year or two but may be longer, depending on how far the economy is initially from conditions of maximum employment and price stability. (Bernanke 2011, pp. 4–5)

For another example of the use of the projections, in June 2010 the FOMC's projections for underlying inflation were below the mandate-consistent level, and its projections for unemployment were above the estimate of the sustainable unemployment rate. Indeed, with reference to these circumstances, Chair Bernanke (2010) concluded, in a speech shortly before the FOMC announced QE2:

> Given the Committee's objectives, there would appear—all else equal—to be a case for further action.

However, the median projections of the federal funds rate, inflation, and unemployment in the SEP are obviously conceptually different from the forecast-targeting policy rate path and forecasts of inflation and unemployment discussed previously. Importantly, the projections in the SEP are not the result of a joint FOMC decision. The median projections reported are the medians of the modal projections of each individual FOMC participant (that is, voter or nonvoter) rather than of each individual member (voter). It thus gives equal weight to voters and nonvoters. Furthermore, all voters need not have the same weight in the decision; in particular, the Chair has more weight than others in the decision.

Also, the median projections of the federal funds rate, inflation, and unemployment are inconsistent, in the sense that they are not the projections of a median participant. Instead they may consist of a combination of projections of different participants, combinations that are likely to vary across the federal funds rate, inflation,

and unemployment. In addition, the participants may have different models of the economy and the transmission of monetary policy. Thus, the median projection of the federal funds rate is generally not consistent with the median forecasts of inflation and unemployment. Publishing participants' initials with the projections would provide more information and moderate some of these problems.

However, even if the SEP is conceptually different from the forecast-targeting policy rate path and forecasts of inflation and unemployment, it is not clear how quantitatively different they are from a joint FOMC decision. Majority voting about paths in a committee may result in medians consisting of sections from different committee members, but it is not clear whether this would be problem of quantitative importance.[38] But it is clear that the SEP is more of a snapshot of the views of the FOMC participants and that the medians of the SEP are not a conscious joint decision by the FOMC.

A decisionmaking process whereby the FOMC arrives at an explicitly joint policy rate path and corresponding inflation and unemployment forecasts would be more consistent with forecast targeting. The FOMC has undertaken some experiments in constructing a consensus policy rate path and forecasts of inflation and unemployment. They are discussed in some detail under the heading "Experimental Consensus Forecast" in the October 2012 transcripts (FOMC 2012, pp. 201–79). There were several difficulties noted about constructing consensus forecasts. According to a summary of a staff memo about the experiment presented at the meeting, among these difficulties, reaching a consensus on the appropriate medium- and longer-term policy path could be extremely difficult. One reason was that the policymaking environment was unusually complex, with both unconventional portfolio actions and forward guidance being important policy tools. Another reason was that participants who could agree on the appropriate policy action to be taken at a given meeting might nevertheless disagree about the

[38]Svensson (2007b) discusses majority voting on forecast paths and argues that they are completely feasible and already occurring in a few central banks. For example, the nine-member Monetary Policy Committee of Bank of England makes decisions on the quarterly forecast paths of inflation, unemployment, and GDP growth three years out. A 12-member FOMC should be able to do the same and include a policy rate path as well.

appropriate stance of policy further out in the future. Some, but not all, such disagreement might be because participants would disagree about the likely future evolution of asset purchases. But it was noted that even in normal times, the FOMC has typically only described its policy decision in terms of the change in the funds rate agreed to at a particular meeting, not the anticipated future path of policy.

There were also some production-related challenges. Because the Committee's policy decisions are not known in advance of the meeting, it would not be possible to guarantee the production of a forecast that incorporates the Committee's policy decision in time for the Chair's press conference. A required delay would reduce the usefulness of the forecast for communication purposes. The participants were also not clear about how they should determine and express whether they support the proposed consensus outlook.

In view of these difficulties, the FOMC abandoned the consensus forecast exercise at the time—perhaps not permanently—and instead focused on improvements on the SEP.

It is obvious that a decision process that includes reaching a consensus on the policy rate path faces difficulties when the FOMC also decides on balance sheet policies and thus has several policy instruments. However, when the balance sheet reduction is set on autopilot and the FOMC has the federal funds rate as its one policy instrument, perhaps such a decisionmaking process is nevertheless possible and can be followed, perhaps with some iterations between the Board and the Banks. From a production point of view, having the press conference the next day—as is done at the Riksbank—may help with the production problems.

Nevertheless, forecast targeting has arguably to a considerable extent already been practiced with the current median projections in the SEP, as the quotes above from Bernanke (2010, 2011, 2015b) suggest. Furthermore, the description in the section "Systemic Monetary Policy in Practice" of Federal Reserve Board (2019, p. 39) of how the FOMC currently conducts its policy is very similar to the description of forecast targeting here. In addition, the FOMC can to some extent be held accountable in real time with the current SEP. With some reservations due to the problems mentioned, it is possible to compare the median projection of the federal funds rate with market expectations and the median projections of inflation and unemployment with, respectively, the 2 percent target and

the FOMC's estimate of the long-run sustainable rate and assess whether the FOMC is best fulfilling its mandate.

6. The New Monetary Policy Strategy

In August 2020, the Federal Reserve announced a revision of its monetary policy strategy and released a new "Statement on Longer-Run Goals and Monetary Policy Strategy" (FOMC 2020), a result of the review of its strategy that it had initiated in 2019 (Powell 2020). With regard to the *maximum-employment mandate*, the FOMC now seeks over time to mitigate "shortfalls" of employment from its maximum level, not "deviations." This means that a low unemployment rate by itself, unless accompanied by signs of unwanted increases in inflation, will not justify a policy tightening. Focusing on shortfalls of employment instead of deviations introduces an asymmetry in the maximum-employment mandate, and the statement drops previous language about "a balanced approach."

With regard to the *price-stability mandate*, the FOMC now "seeks to achieve inflation that averages 2 percent over time." This implies that if inflation has been running persistently below 2 percent, the FOMC would likely aim to achieve inflation "moderately above 2 percent for some time." The Federal Reserve has thus adopted an explicit "makeup" strategy. As explained by Powell (2020), Clarida (2020), and Brainard (2020), this introduces a strategy of "flexible average inflation targeting." It is also made clear that it would not be appropriate to implement this strategy by using a mechanical Taylor-type instrument rule (Clarida 2020; Brainard 2020).[39]

What changes to forecast targeting as described above do these changes in the Federal Reserve's strategy imply? Svensson (2020) examined alternative monetary policy strategies for the Federal Reserve and concluded that flexible average inflation targeting—implemented by forecast targeting rather than by a mechanical Taylor-type rule—offers some advantages over the other strategies.

[39]Such reservations were also expressed by FOMC participants in the discussion of average inflation targeting at the September 2019 FOMC meeting (FOMC 2019a).

For average inflation targeting, the loss function (2) is then simply modified to include the term $(\bar{\pi}_t - \pi^*)^2$, where $\bar{\pi}_t$ denotes a measure of the average inflation rate, $\bar{\pi}_t$.[40] In order to avoid too much variability of the annual inflation rate, the FOMC may want to put some weight on stabilizing both the average inflation rate and the annual inflation rate (Svensson 2020, eq. (17)).[41] Then the forecast loss function (4) can be modified accordingly.

The asymmetry introduced by mitigating only shortfalls of employment from its maximum level can be literally interpreted as replacing the quadratic unemployment term in the loss function (2) with $(u_t - u^*)^2$ if $u_t \geqslant u^*$ and with 0 if $u_t < u^*$. This does not assess any loss at all to a shortfall of the unemployment rate from its minimum long-run sustainable rate. Such an asymmetric loss function implies that the conditions for certainty equivalence are no longer satisfied, and mean forecasts are no longer sufficient statistics for optimal monetary policy (section 3.1). This may cause considerable difficulties, both technical and practical, in the conduct of monetary policy.[42]

However, if the shortfalls of employment from its maximum level is *operationally* interpreted as referring to shortfalls of the (mean) *forecast* of employment, this can be interpreted as an asymmetric *forecast* loss function. That is, the forecast loss function (4) is modified such that the quadratic unemployment-forecast term is replaced by $(u_{t+\tau,t} - u^*)^2$ if $u_{t+\tau,t} \geq u^*$ and by 0 if $u_{t+\tau,t} < u^*$.

[40]Svensson (2020) uses an averaging period of five years as an example. The FOMC retains some flexibility by not (yet) specifying the averaging period.

[41]This is consistent with the inclusion of the word "moderately" in the quote "moderately above 2 percent for some time."

[42]Such a loss function with an asymmetry from output deviations has been used by Cukierman and Gerlach (2003). The linear-exponential (linex) function allows asymmetry in a more flexible way and has been used in monetary policy applications with an asymmetry for inflation deviations by, for example, Chadka and Schellekens (1999), Nobay and Peel (2003), and Ruge-Murcia (2003) and, with an asymmetry in both inflation and output deviations, by Surico (2007). Monetary policy with asymmetric loss functions generally results in an inflation bias. The absence of certainty equivalence requires quite complex analysis. Erceg, Kiley, and Tetlow (2014) provide a perfect-foresight simulation of optimal monetary policy for a linex loss function where the cost of a below-target outcome is considerably lower for both unemployment and inflation than of a corresponding above-target outcome.

Given this modification of the forecast loss function—including the average-inflation modification mentioned above—it is straightforward to accordingly modify the intertemporal forecast loss functions (5), and the mean squared gaps for inflation (7) and unemployment (8). The concept of forecasts "looking good"—best fulfilling the Federal Reserve's mandate—then means best stabilizing average inflation around its target and mitigating overshoots of the unemployment forecast from its long-run sustainable rate. Then forecast targeting can directly be applied to the new monetary policy strategy.

7. Conclusions

If the FOMC seeks to fulfill its mandate of maximum employment and price stability, while also being held accountable for fulfilling that mandate, forecast targeting is likely to dominate a Taylor-type rule. Forecast targeting means selecting a policy rate and policy rate path so that the forecasts of inflation and employment (or unemployment) "look good," in the sense of best stabilizing inflation around the Federal Reserve's target of 2 percent and employment around its maximum level. The justification may involve demonstrations that other policy rate paths would lead to worse mandate fulfillment. Publication and justification may contribute to making the policy rate path and the forecasts credible with the financial market participants and other economic agents and thereby more effectively implement the FOMC's policy. Importantly, with such information made public, external observers and experts can review FOMC policy, both in real time and after the fact, that is, after the outcomes for inflation and employment have been observed. This way the FOMC can be held accountable for fulfilling its mandate. In contrast to simple policy rules that rely on very partial information in a rigid way, such as a Taylor-type rule, forecast targeting allows all relevant information to be taken into account and has the flexibility and robustness to adapt to new circumstances.

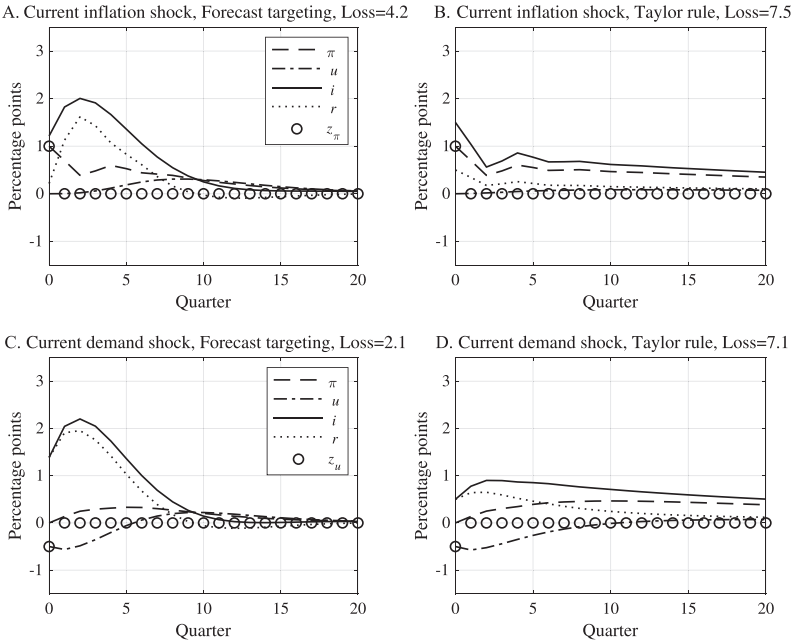
Furthermore, the FOMC is already practicing forecast targeting to a considerable extent, with the publication in the Summary of Economic Projections of the participants' projections of the federal funds rate, inflation, and unemployment. Although these projections

are not a joint decision of the FOMC, it is not clear how quantitatively different they are from a joint decision. They have already been used by the Chair to explain and justify policy decisions, and they are also used by external observers to some extent to hold the FOMC accountable for fulfilling its mandate. Making the projections a joint decision by the FOMC would make them more suitable for explaining and justifying the decision and for holding the FOMC accountable for fulfilling its mandate.

The Federal Reserve's new monetary policy strategy announced in August 2020 implies flexible *average* inflation targeting together with an asymmetry in the form of less or no cost of employment exceeding the maximum sustainable level. If the asymmetry is operationally interpreted as referring to employment *forecasts*, the appropriate modification of the forecast loss function is straightforward, and forecast targeting can be directly applied to the new strategy.

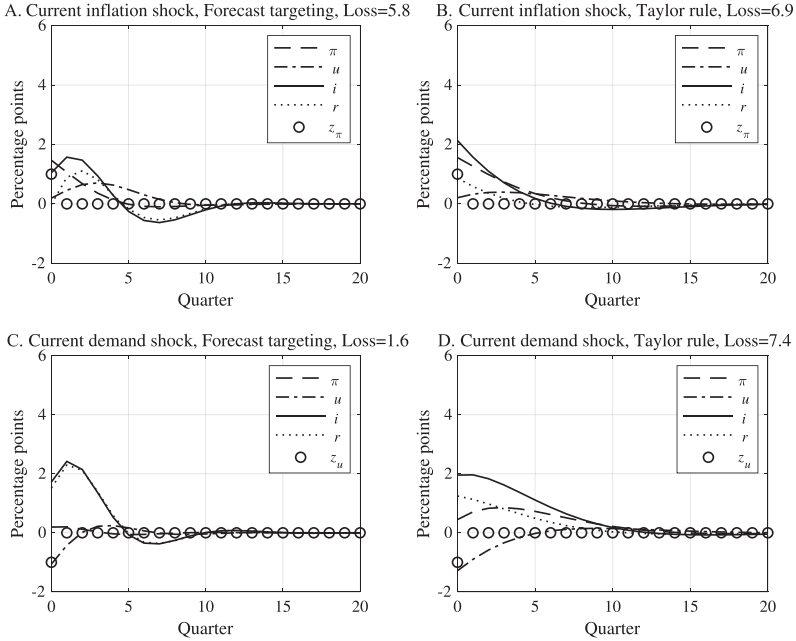
Appendix. The Backward- and Forward-Looking Model: Impulse Responses

Figure A.1. Forecast Targeting and the Taylor Rule: Backward-Looking Model, Impulse Responses



Notes: The period loss function is $L_t = \pi_t^2 + u_t^2 + 0.2(i_t - i_{t-1})^2$. The Taylor rule is $i_t = 1.5\,\pi_t - u_t$. The real interest rate that affects the unemployment gap in the Rudebusch and Svensson (1999) model is $r_t = i_t - \pi_t$, which is plotted in the figure. The intertemporal loss reported excludes the interest rate term, which is small. See Svensson (2005) for details. In the top two panels, there is an inflation shock of 1 percentage point in quarter 0. In the bottom two panels, there is an unemployment shock of -0.5 percentage point (a GDP shock of 1 percentage point) in quarter 0.

Figure A.2. Forecast Targeting and the Taylor Rule: Forward-Looking Model, Impulse Responses



Notes: Forecast targeting minimizes the loss function $L_t = \pi_t^2 + u_t^2 + 0.2(i_t - i_{t-1})^2$. The Taylor rule is $i_t = 1.5\,\pi_t - u_t$. The real interest rate that affects the unemployment gap in the Lindé (2005) model is $r_t = i_t - \pi_{t+1|t}$, which is plotted in the figure. The intertemporal loss reported excludes the interest rate term, which is small. See Svensson (2005) for details. In the top two panels, there is an inflation shock of 1 percentage point in quarter 0. In the bottom two panels, there is an unemployment shock of -1 percentage point (a GDP shock of 2 percentage points) in quarter 0.

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The Aggregate and Country-Specific Effectiveness of ECB Policy: Evidence from an External Instruments VAR Approach*

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This paper studies the transmission of ECB policy, both at the aggregate euro-area level and the country level. We estimate a VAR model for the euro area in which monetary policy shocks are identified using an external instrument that reflects unexpected changes in the policy stance. For that purpose, we use changes in German bunds at meeting days of the Governing Council and selected intermeeting announcements. We also decompose policy shocks into pure policy surprises and information shocks. The resulting impulse responses are robust with respect to the choice of the instrument. Expansionary monetary policy affects prices and real activity but remains ineffective in pushing credit and stock markets. We show that pure policy shocks, i.e., shocks net of the new information revealed on meeting days, also have a significant effect on credit and stock prices. The identified monetary policy shock is then put into country-specific local projections in order to derive country-specific impulse responses. The transmission is heterogeneous across member countries with credit and financial markets being unevenly affected by monetary policy.

JEL Codes: E52, E44, E32.

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1. Introduction

In the aftermath of the 2008–09 financial crisis and the subsequent European debt crisis, the European Central Bank (ECB) adopted a series of unconventional policy measures. More precisely, with short-term interest rates at the effective lower bound, the ECB used unconventional monetary policy such as the Asset Purchase Programme (APP) and the targeted longer-term refinancing operations (TLTROs) to provide additional stimulus. Given the persistently low level of inflation and the sluggish recovery despite several years of expansionary monetary policy, the assessment of ECB policy is controversial. Only very recently, the recovery in the euro area gained momentum. Since 2008, analyzing monetary policy has become more difficult, as the overall policy stance is no longer appropriately summarized by the short-term policy rate. In fact, the ECB uses several instruments at the same time. Moreover, with a large share of monetary policy being transmitted through asset markets and this share becoming larger over the recent years, identifying monetary policy shocks has become more difficult. The traditional triangular identification scheme applied to vector autoregressive (VAR) models that imposes restrictions on the contemporaneous interaction among the variables is not suitable with financial data. Sign restrictions, a popular alternative to the Cholesky ordering, require imposing more or less controversial restrictions onto the dynamic interaction.

In this paper, we study the monetary policy transmission in the euro zone, both at the aggregate euro-area level and the disaggregated country level. For that purpose, we use an external instruments VAR approach to identify an ECB policy shock. The external instruments approach, which has recently been made popular by the work of Stock and Watson (2012), Mertens and Ravn (2013), and Gertler and Karadi (2015) identifies the simultaneous dynamics of monetary policy and asset prices with the help of the behavior of an instrument on central bank meeting days. The assumption is that around an ECB announcement, the instrument reflects only the policy surprise, which is orthogonal to other potential shocks driving the VAR system.

Based on the identified policy shock, we make the following contributions: First, we provide evidence on the effects of a monetary policy shock at the aggregate euro-area level for a full 2002–16 and a

post-crisis sample. Expansionary monetary policy affects consumer prices and real activity and leads to a depreciation of the euro in real terms. While the shock also compresses the corporate bond spread, monetary policy remains ineffective in pushing credit and stock markets.

Second, we take account of the recent literature on information shocks, highlighting the fact that policy decisions by central banks also reveal information about the central bank's assessment of the economy (Jarocinski and Karadi 2018, Miranda-Agrippino and Ricco 2018, and Nakamura and Steinsson 2018). Under incomplete information, these information shocks are distinct from the pure monetary policy component of shocks. We use a principal component analysis to decompose the monetary policy shock into an information shock and a pure policy shock and find plausible impulse responses to both shocks. Based on this decomposition, we are able to show that the baseline results remain robust when we exclude the information revealed on meeting days from the monetary policy surprise.

Third, we use the identified euro-area policy shock to estimate several country-specific impulse response functions from local projections (Jordà 2005). This provides us with the effects of the common monetary policy on individual countries and excludes the feedback from the country level to ECB policy. The assumption is that the ECB is, in line with its mandate, directing policy to the euro-area aggregate, not to specific countries. The results show homogenous cross-country responses for consumer prices and industrial production but heterogeneity in the effects of monetary policy across members on unemployment, credit, and the stock market. In several countries the transmission through equity prices and through the banking system in terms of bank lending is severely dampened. The cross-country heterogeneity in the effects of bank lending and stock markets reflects the insignificant responses of both variables at the euro-area level.

Our project connects several strands of the recent literature: Hachula, Piffer, and Rieth (2019) and Andrade et al. (2016) also use an external instruments approach to estimate euro-area VAR models. However, their focus is different. The first paper estimates the effects of monetary policy shocks on fiscal policy variables in the euro area and studies whether fiscal discipline deteriorates after a monetary policy easing. The authors indeed find an increase in

public expenditure after an expansionary policy shock. Andrade et al. (2016) focus on the ECB's Asset Purchase Programme, implemented since January 2015.[1] Two other recent papers, namely Cesa-Bianchi, Thwaites, and Vicondoa (2016) and Ha (2016), use the external instruments approach for shock identification in an open-economy VAR model and put the shock series into local projections.[2]

Furthermore, Wieladek and Pascual (2016) use a Bayesian VAR model with a battery of alternative identification schemes to study the euro area in 2012–16. Counterfactuals for the euro-area and for the country level show that monetary policy has a very large effect. Since January 2015, it has led to real gross domestic product (GDP) being 1.3 percent higher than in the absence of quantitative easing (QE). The same policy has benefited Spain the most and Italy the least. Boeckx, Dossche, and Peersman (2017) use a sign-restricted VAR model to study the effects of unconventional monetary policy shocks that drive up the ECB's balance sheet.[3] Based on a Bayesian VAR model, Mandler, Scharnagl, and Volz (2016) provide evidence for heterogeneous ECB policy transmission across the four largest economies of the euro area.

While the previously mentioned papers work with monthly or quarterly data, Fratzscher, Lo Duca, and Straub (2016) use daily data to study the responses of a broad range of asset prices to ECB announcements prior to 2013. They find that unconventional policy boosts asset prices and spills over to other economies' equity markets but not to other bond markets. The work by Burriel and Galesi (2018) also focuses on euro-area and country-specific effects of policy. The authors estimate a global VAR model for the euro area which allows for spillovers among euro-area countries. They find these intra-EMU spillovers to be sizable. In addition, they document a large heterogeneity of cross-country effects of monetary policy shocks.

[1] A very useful survey of the transmission channel of unconventional ECB policy is provided by Fiedler et al. (2016).

[2] Altavilla, Darracq Paries, and Nicoletti (2015) construct an indicator of credit supply tightening in the euro area and include it as an external instrument in a VAR model.

[3] Hristov et al. (2014), Altavilla, Giannone, and Lenza (2016), and De Santis (2016) provide additional evidence on selected ECB programs, such as the Outright Monetary Transactions program and the Asset Purchase Programme.

This paper proceeds as follows: section 2 outlines the VAR model with an external instrument, which is our benchmark model, as well as the data used. The section also discusses our findings for the aggregate euro area, presents results for variables that describe specific transmission channels of monetary policy, and also introduces the decomposition of policy surprises into pure policy shocks and information shocks. Section 3 introduces the local projections approach and discusses the country-specific results. Section 4 draws on these findings and discusses policy implications.

2. A Euro-Area VAR Model with External Instruments

2.1 Methodology

In this subsection, we describe how we combine the conventional VAR methodology with the event-study approach. We build upon the methodology of Stock and Watson (2012), Mertens and Ravn (2013), and Gertler and Karadi (2015) in order to overcome the problems of endogeneity without imposing sign or zero restrictions. The endogeneity issue is particularly relevant for financial variables, which are supposed to react instantly to a monetary policy shock. In line with, e.g., Gambetti and Musso (2017), we expect unconventional monetary policy to influence financial variables. Therefore, a Cholesky ordering can potentially provide misleading results. It is also hard to argue in favor of sign or zero restrictions. Upon imposing restrictions, presumptions about the behavior of the included variables have to be made. This is problematic in the case of unconventional monetary policy, where we know very little about its transmission. However, under the assumption that an accurate instrument can be found, we are able to capture the transmission of the complete set of monetary policy tools.

Our goal is to derive the structural VAR model according to equation (1):

$$S^{-1}Y_t = C + \sum_{j=1}^{p} B_j Y_{t-j} + \sum_{k=0}^{q} D_k X_{t-k} + u_t. \tag{1}$$

Hereby, Y_t represents the endogenous and X_t the set of exogenous variables at time t. While C captures constants, the matrices B_j and D_k contain the coefficients on the lags of the endogenous and exogenous variables up to lag length j and k, respectively. The simultaneous effect of one endogenous variable to another is captured by S^{-1} and u_t stands for the vector of error terms.

Due to the endogenous nature of the variables in Y_t, we are not able to solve the structural VAR uniquely. Hence, we first estimate the reduced-form VAR, which results after multiplying each side of equation (1) by S:

$$Y_t = S \cdot C + \sum_{j=1}^{p} S \cdot B_j Y_{t-j} + \sum_{k=0}^{q} S \cdot D_k X_{t-k} + \varepsilon_t. \tag{2}$$

The reduced-form innovations are then given by equation (3):

$$\varepsilon_t = S \cdot u_t. \tag{3}$$

Here S is a square matrix with the dimension equal to the number of endogenous variables. The i-th column in S captures the response of the vector of reduced-form innovations, ε_t, to an increase in the i-th element of the matrix of structural shocks u_t. As we are only interested in the responses to a structural monetary policy shock u_t^{MP}, we just have to identify the column s in S that captures the impact of u_t^{MP} on the vector ε_t. Now let ε_t^{MP} be the reduced-form innovation of the monetary policy equation and s^{MP} be the element of s that describes its response to the structural shock, u_t^{MP}, such that equation (4) holds:

$$\varepsilon_t^{MP} = s^{MP} \cdot u_t^{MP}. \tag{4}$$

Accordingly, ε_t^q and s^q are reduced-form error terms and the respective elements in s that correspond to other variables:

$$\varepsilon_t^q = s^q \cdot u_t^{MP}. \tag{5}$$

Solving for u_t^{MP} in equations (4) and (5) leads to

$$u_t^{MP} = \frac{\varepsilon_t^{MP}}{s^{MP}} = \frac{\varepsilon_t^q}{s^q}, \tag{6}$$

which can be rearranged to

$$\varepsilon_t^q = \frac{s^q}{s^{MP}} \varepsilon_t^{MP}. \tag{7}$$

Finally, with the reduced-form error terms as both the dependent and the explanatory variable, respectively, an estimate for $\frac{s^q}{s^{MP}}$ can be found. In order to overcome the possible endogeneity of ε_t^q and ε_t^{MP}, we apply a two-stage least-squares approach. From the first stage we receive $\widehat{\varepsilon_t^{MP}}$ as an estimate that only captures changes in monetary policy that do not stem from a simultaneous change in ε_t^q. In the second stage, we then simply run the following ordinary least-squares (OLS) regression:

$$\varepsilon_t^q = \frac{s^q}{s^{MP}} \widehat{\varepsilon_t^{MP}} + \xi_t. \tag{8}$$

Given these estimates and the variance-covariance matrix of the reduced-form VAR model, we are able to uniquely identify all components of s. The crucial point in this framework is to find an accurate instrument Z_t which is, by definition, correlated with ε_t^{MP} but orthogonal to ε_t^q.

2.2 Data

For our baseline euro-wide model, the vector of the endogenous variables consists of the log of industrial production (excluding construction), the log of the Harmonised Index of Consumer Prices (HICP), a corporate bond spread, and the (shadow) short rate.[4] Following Sims (1992), we further add (the log of) oil prices as an exogenous variable in order to avoid the price puzzle.

Prior to the financial crisis, the ECB conducted open market operations in order to move the key policy rate. With the zero lower bound (ZLB) and the introduction of unconventional monetary

[4]In the baseline model, the corporate bond spread is the spread between the yield on BBB-rated and AA-rated bonds. However, spreads between corporate bonds with other ratings lead to similar results.

policy, the ECB extended its policy toolkit. It is for this reason that we use the (shadow) short rate provided by Wu and Xia (2016) for the interval available (i.e., from 2004:M9 until the end of our sample) as the measure of the monetary policy.[5] Until 2004:M8 the euro overnight index average (EONIA) rate represents the monetary policy stance, which we receive from Thomson Reuters Datastream. We generally draw on seasonally adjusted data for the changing composition of the European Economic and Monetary Union (EMU). Financial variables that are not expected to contain seasonal patterns are not adjusted. A complete list of all variables, their adjustment, and their sources can be found in table A.1 in the appendix. The sample consists of monthly data from 2002:M1 until 2016:M10.[6]

We include six lags as suggested by the Akaike information criterion and the final prediction error. However, as outlined below, choices of other lag lengths lead to similar results.

After estimating the baseline four-variable model, we add a fifth variable to our baseline model to shed light on several aspects of the transmission process. This fifth variable is taken from the following list of variables: euro-area government bond yields, the unemployment rate, the log of the real exchange rate, the log of the Euro Stoxx 50, the log of the rent component of the HICP, the log of the loan volume granted by financial institutions, and the net percentage change of credit standards and credit demand, both obtained from the Bank Lending Survey.[7]

[5]We also apply the shadow rates provided by Krippner (2012) and Lemke and Vladu (2017) as well as the zero-coupon one-year German government bond rate. As will be outlined below, the results from the three different short rates are indeed complementary.

[6]Data from the ECB's Bank Lending Survey as well as loan data are only available from 2003 onward. Hence, a shorter sample size is used for models containing these variables.

[7]Within the Bank Lending Survey the banks answer whether they tightened lending standards "considerably" or "somewhat," eased "somewhat" or "considerably," or left the standards unchanged. The net percentage change is the difference in the percentage of banks that tightened their lending standards (either "somewhat" or "considerably") and the share of banks that eased them. Accordingly, the net percentage change in the credit demand is the share of banks that expect an increase in the demand for loans (either "considerably" or "somewhat") minus the share that expect a decrease in the demand.

2.3 Choosing an Instrument

The choice of the instrument deserves special attention. We use changes in the German 10-year government bond yield on meeting days and a small number of other selected dates as the instrument.

The rationale behind the use of daily changes, rather than intraday data, lies in the timing of ECB communication on meeting days of the Governing Council. The press release at 13:45 CET on every meeting day is followed by a press conference at 14:30 CET. Since our instrument has to capture the market response to the press conference as well, we cannot apply the widely used 30-minute window.[8] The data of all external instruments stem from Thomson Reuters Eikon.

The financial crisis and the subsequent European debt crisis opened up an interest rate spread between government bonds of various euro-area countries. While the yield on German government bonds serves as a risk-free rate throughout the entire sample, the status of government bonds of other countries switches from a risk-free to an exposed asset. With the choice of German government bonds, we avoid the issue of a structural break within our instrument variable. Furthermore, we consider 10-year bonds since our applied instrument also has to reflect changes in investors' expectations through unconventional monetary policy measures such as forward guidance.

Our identification method rests on the efficient market hypothesis (EMH). The EMH states that movements in asset prices only appear if new information is received. Thus, under the assumption that news other than the monetary policy decisions on the meeting days and the selected special events are white noise, the changes in German bond yields on these days represent changes in the monetary policy stance. For example, an increase in the German bond yield on these days, i.e., a positive surprise component, reflects a monetary tightening.

[8] In fact, Gürkaynak, Sack, and Swanson (2005) find that daily changes in federal fund futures on Federal Open Market Committee meeting days are akin to changes in a 30-minute window around the release in the time span from 1994 to 2004. Thus, they conclude that "the surprise component of monetary policy announcements can be measured very well using daily data."

With the adoption of unconventional policies, important news about monetary policy also emerged on nonmeeting days. Hence, we supplement the set of meeting days by three additional events. These are the announcement of the two tranches of the Securities Markets Programme (SMP) on May 10, 2010 and August 7, 2011, respectively, as well as President Draghi's "whatever-it-takes" speech on July 26, 2012. The monthly series for our instrument consists of the change in German yields on these specific days—that is, if the Governing Council meets on one Thursday in a given month, the yield change on this day is used as the monthly entry in the instrument series. If there is both a Governing Council meeting and one of the additional events in a given month, we sum up the yield changes on these two days in order to get an estimate for the surprise component of that month.

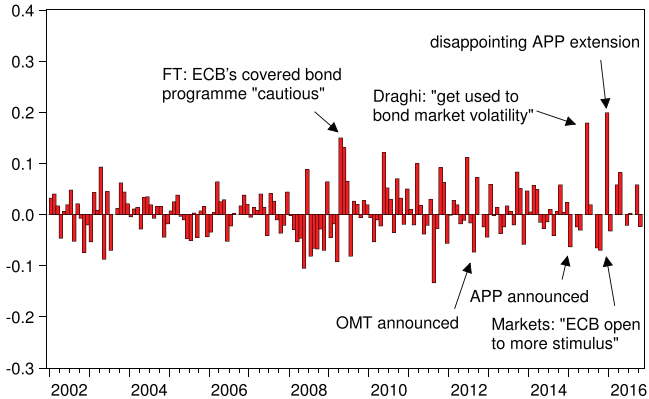
This measure for the monetary policy stance has several advantages. First, the surprise component serves as a consistent measure for the entire monetary policy toolkit. With the ECB adopting unconventional policies, it extended its set of policy instruments. By having one measure reflecting the entire set of policy instruments, we do not face the problem of disentangling the effects of each instrument, which is particularly challenging as those have been used simultaneously.

Second, the focus on market reactions allows us to directly measure the unanticipated part of a policy change. This is better suited for identifying a policy shock, as according to the EMH only those should influence asset prices. For example, an increase in the interest rate that is lower than expected is recognized as an expansionary monetary policy in the view of market participants. Finally, the external instruments approach clearly defines an unexpected monetary policy shock, which is the starting point of every analysis within the VAR model.

The series of the surprise component from 2002 until 2016 is plotted in figure 1. As the surprise component fluctuates around zero, it can be concluded that there is no systematic bias in the market expectations.[9] The largest swings are found after the financial crisis in 2007. President Draghi's remark "get used to market volatility"

[9]On a 10 percent significance level, a t-test confirms that the mean of the surprise component is not different from zero.

Figure 1. Monetary Policy Surprises Obtained from 10-Year Bunds



Notes: Policy surprises are defined as the change in the yield on 10-year German bunds on ECB meeting days and selected other days. This series is used as an external instrument in the VAR identification. The annotation refers to the *Financial Times* (FT), the Outright Monetary Transactions (OMT) program and the Asset Purchase Programme (APP).

on June 2015 and the disappointment about the size of the additional stimulus adopted in December 2015 account for the peaks in the surprise component. In contrast, the announcements of the Outright Monetary Transactions (OMT) program in September 2012 and the APP in January 2015 are reflected in negative surprises. In other words, monetary policy was surprisingly expansionary.

Before we turn to the results of our VAR model, we check if the considered instrument is accurate. First, we test for the information content of the instrument in an event study. We run the regression

$$\Delta y_t^{daily} = \alpha + \beta \cdot \Delta Z_t^{events} + \epsilon_t, \tag{9}$$

where the daily changes in asset prices, y_t^{daily}, are regressed on a constant and the surprise component, i.e., the changes in the German 10-year bond yield Z_t^{events}, using OLS. For this estimation we *only* consider the selected events, i.e., meeting days of the Governing Council and three selected special events, which leaves us with a total of 175 observations. The list of dependent variables consists of the log of the U.S. dollar exchange rate to the euro, the euro-area

Table 1. Monetary Policy Surprises in an Event Study

y_t		Coef.	p-value
(log) Exchange Rate	$\hat{\alpha}$	0.000	0.43
	$\hat{\beta}$	0.071	0.00
EURIBOR Future	$\hat{\alpha}$	−0.006	0.15
	$\hat{\beta}$	0.890	0.00
Corporate Bond Spread	$\hat{\alpha}$	−0.005	0.12
	$\hat{\beta}$	0.192	0.00

Notes: Results from an event-study regression of y_t on policy surprise series with the slope coefficient β and a constant α.

interbank offered rate (EURIBOR) futures rate,[10] and the corporate bond spread.[11]

The results of the regressions are presented in table 1. An unexpected increase in the government bond yield on meeting days, i.e., a surprise tightening, leads to an appreciation of the euro and increases in the EURIBOR future and the corporate bond spread. This suggests that changes in the German bond yield indeed contain information about unexpected changes in the ECB's monetary policy stance.

We further evaluate the properties of the instrument by testing for the occurrence of a weak instruments problem. The explanatory power of the instrument can be examined by regressing the reduced-form VAR residuals of the monetary policy equation on a constant and the external instrument. As described by Li and Zanetti (2016), this equals the first stage in our two-stage least-squares regression from equation (8). For the changes in the German 10-year bond yields, the corresponding F-statistic in the baseline case is 10.44. Following Stock, Wright, and Yogo (2002), a value for the F-statistic lower than 10 indicates a weak instrument issue. With the

[10] At any point in time, we consider the future that is the eighth next to deliver. Note that the first six delivery months are consecutive in time. Given that the subsequent delivery months—namely March, June, September, or December—settle on a quarterly frequency, the delivery of the future that we consider is roughly in one year. Our presented results are robust to other continuation futures.

[11] The corporate bond spread presented here is the spread between A- and BBB-rated bonds.

German bond yields avoiding the weak instrument problem and showing plausible results for the event-study regression, we are confident about our choice of an accurate instrument. Hence, we are able to estimate the impulse responses from our VAR model. The results are discussed below.

2.4 Results

We start by estimating the effect of an expansionary monetary policy that leads to a 25 basis point (bp) drop in the shadow rate. All results are presented as impulse response functions together with a 90 percent confidence interval.

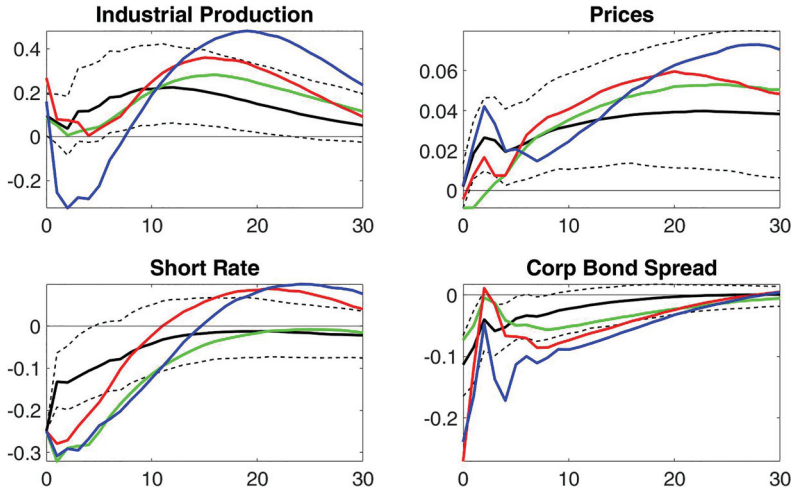
2.4.1 Baseline Model

The results from the baseline VAR model are presented in figure 2. As indicated by the black lines, the responses of industrial production, prices, and the corporate bond spread to an expansionary shock have the expected sign and are statistically significant. As noted above, we circumvent the price puzzle by adding oil prices as an exogenous variable, so that a monetary easing immediately increases prices. The responses of the consumer price index (CPI) and the industrial production index indicate that ECB policy stimulated both inflation and real economic activity. Boeckx, Dossche, and Peersman (2017) find similar results by imposing sign restrictions in a euro-area VAR model. Following Zhu (2013), the corporate bond spread reflects the external finance premium and, hence, the credit channel of monetary policy transmission. We find that spreads narrow immediately upon the monetary easing, which is consistent with the presence of the credit channel.

Further on, we review the accuracy of our outcome by altering the shadow rate and the lag length. In this respect, the green and blue lines in figure 2 show the impulse responses based on the shadow rates of Krippner (2012) and Lemke and Vladu (2017), respectively.[12] The results turn out to be similar. Gertler and Karadi (2015) have used a safe interest rate with a maturity of one year,

[12]For color versions of the figures, see the online version of the paper on the IJCB website (http://www.ijcb.org).

Figure 2. Baseline VAR Model



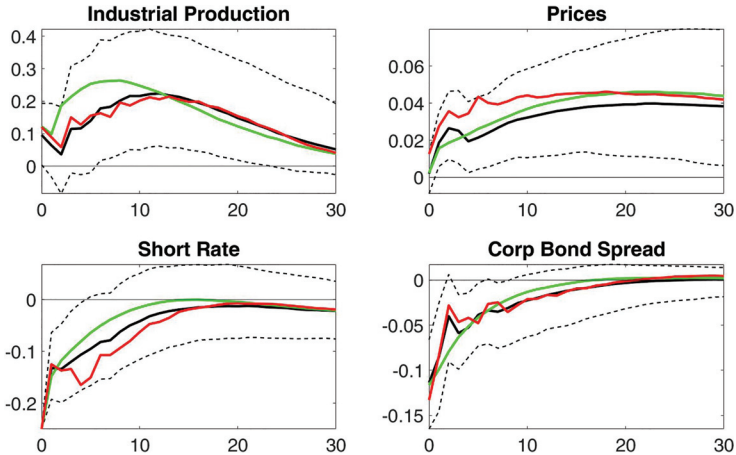
Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model with external instruments and 90 percent confidence band. The black line is the response in the model based on the Wu-Xia (2016) shadow rate, the green line is based on the shadow rate of Krippner (2012), the blue line is based on the shadow rate of Lemke and Vladu (2017), and the red line is estimated based on the one-year German government bond zero-coupon rate.

proxied by the U.S. government bond rate, instead of the shadow rate. For reasons of comparability, we also present results based on the interest rate on a zero-coupon one-year German government bond (red line). Again, the outcomes for industrial production, the price level, and the interest rate are very similar. Only the negative response of the corporate bond spread is more pronounced. Since corporate financing is more dependent on long-term credit conditions, this finding does not come as a surprise. However, one has to keep in mind that, in contrast to shadow rates, the short-term government bond rate hits the zero lower bound during the financial crises. As displayed in figure 3, altering the lag length does not change the results qualitatively.

2.4.2 Cholesky Identification

For a comparison, we apply a Cholesky identification instead of the external instruments approach. The implied ordering of the

Figure 3. Baseline VAR Model: Alternative Lag Structure



Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model with external instruments and 90 percent confidence band. The black line represents the baseline model with a lag length of four. The green (red) line is the impulse response for the model with two (six) lags.

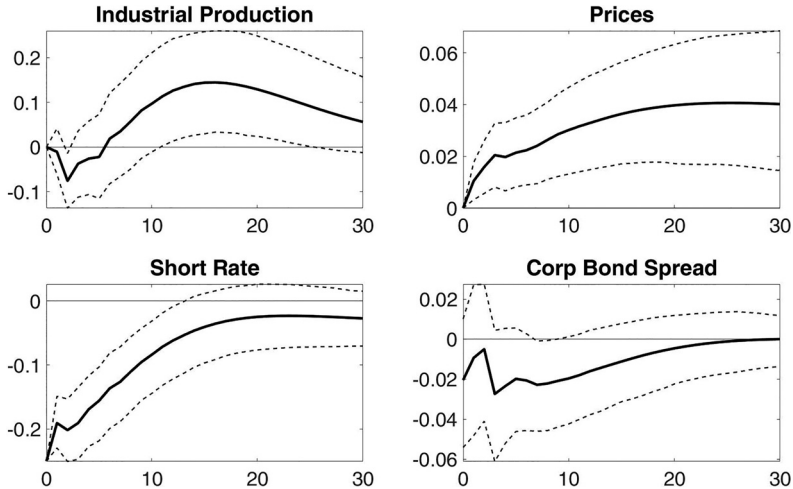
variables is the following: log of industrial production, log of consumer prices, the shadow short rate, and the corporate bond spread. The restriction imposed implies that monetary policy affects the spread contemporaneously, but all other variables with a time lag of one month.

The results are shown in figure 4. While prices and industrial production exhibit responses which are very similar to the baseline findings, the corporate bond spread does not react significantly. This might be the result of the endogenous nature of both the shadow rate and the bond spread, which is not adequately captured by the Cholesky identification. This also lends support to the external instruments approach that we use for identification in our baseline model.

2.4.3 Extending the Baseline Model with Other Real and Nominal Variables

We now turn to the responses of additional variables which were not included in our baseline model. As outlined in the previous

Figure 4. Baseline VAR Model: Cholesky Identification

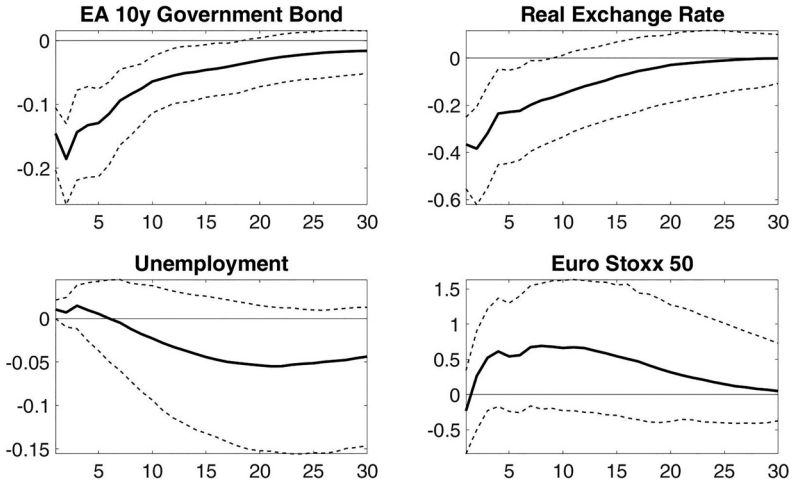


Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model identified recursively and 90 percent confidence band.

subsection, we add one variable at a time as a fifth variable to our model. To save space, we only report the impulse response for the fifth variable.[13] Figure 5 shows the results for euro-area government bond yields, the real exchange rate, unemployment, and the Euro Stoxx 50. Bond yields immediately fall after a monetary easing. Furthermore, the instant depreciation of the euro indicates the existence of the exchange rate channel. Surprisingly, the increase in industrial production found before is not accompanied by a significant decrease in the unemployment rate. Though the sign of the unemployment response is negative, on a 10 percent confidence level, it cannot be ruled out that its response is actually zero. One explanation for the modest decrease in unemployment might be the heterogeneity of business cycles in the euro area. Our results might reflect that, since

[13]Throughout the different VAR models the responses of the four variables in the baseline model remain qualitatively unchanged.

Figure 5. Alternative Fifth Variable: Additional Real and Nominal Variables



Notes: Responses of alternative choices for the fifth variable to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model with external instruments and 90 percent confidence band.

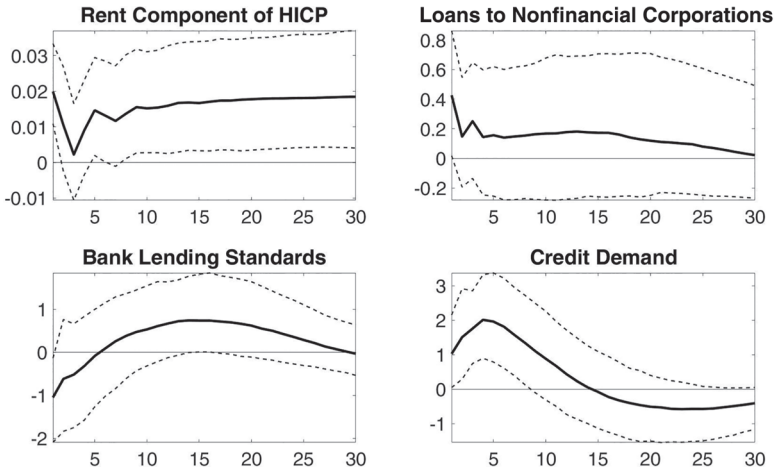
the European debt crisis, unemployment in core and periphery countries respond differently to a monetary policy shock. This hypothesis is supported by our country-specific results presented below.

Although the Euro Stoxx 50 has the expected positive sign, its response turns out to be insignificant. Hence, for the entire time span, we do not find evidence for a policy transmission through the stock market.[14] At a first glance, this seems surprising, as expansionary monetary policy leads to a bull market from a theoretical point of view.[15] However, as outlined in, e.g., Jarocinski and Karadi (2018), Nakamura and Steinsson (2018), and Romer and Romer (2000), a

[14]The reaction of the MSCI Euro Index is virtually identical to the one from the Euro Stoxx 50. These results, along with impulse responses from policy uncertainty, the VSTOXX, and the monetary base, are available on request.

[15]Indeed, Gambetti and Musso (2017) find evidence that the ECB's Asset Purchase Programme increased stock prices.

Figure 6. Alternative Fifth Variable: Credit Market

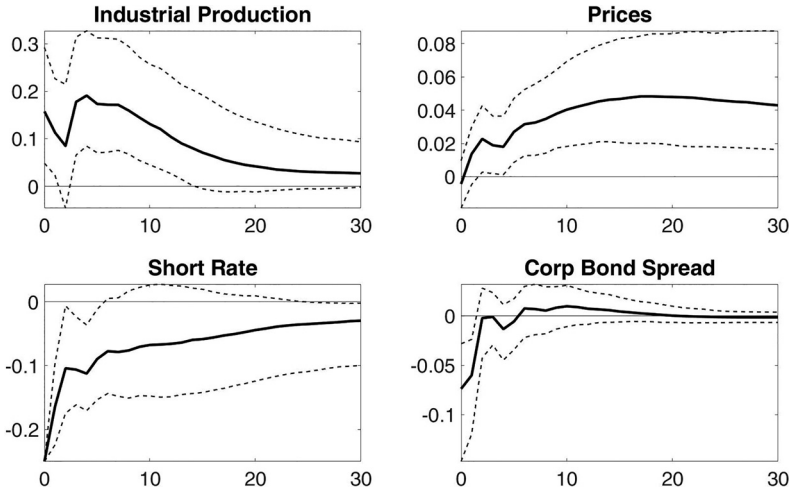


Notes: Responses of alternative choices for the fifth variable to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model with external instruments and 90 percent confidence band.

monetary policy shock also contains information about the policymakers' perceptions of the economic situation. Under the assumption that market participants value this information, the responses of (financial market) variables are also driven by the information component. While a decrease in the short rate due to monetary easing is expected to increase stock prices, a decrease in the policy rate due to weaker economic fundamentals potentially leads to a reduction in stock prices. Below, we follow Jarocinski and Karadi (2018) and further disentangle the information shock from the pure monetary policy shock. Indeed, we find that a pure monetary policy shock (information shock) increases (decreases) the Euro Stoxx 50.

According to figure 6, a monetary easing increases the rent component of the HICP, which serves as a monthly proxy for house prices. A monetary expansion relaxes bank lending standards, thus supporting the existence of a risk-taking channel. The demand for credit increases. A significant reaction in both bank lending and credit demand is also found by Ciccarelli, Maddaloni, and Peydro (2015). Furthermore, we find that the total loan volume to nonfinancial institutions increases.

Figure 7. Baseline VAR Model: 2008:M10–2016:M10



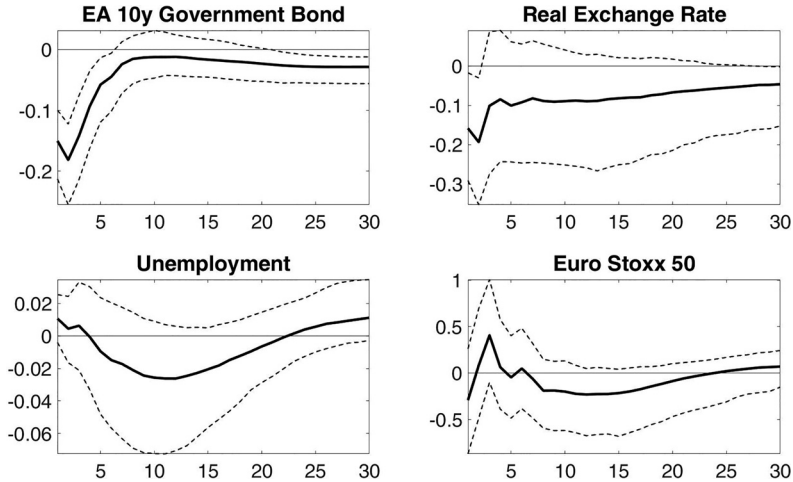
Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model estimated over the post-crisis sample with external instruments and 90 percent confidence band.

2.4.4 The Post-2008 Sample

In order to address the question of how unconventional monetary policy is transmitted, we present evidence from the crisis period only. We interpret the sharp decrease in the ECB's key interest rate as the beginning of the era of unconventional monetary policy. The results based on a sample from 2008:M10 until 2016:M10 are shown in figures 7 and 8. With the shorter time span, we reduce our lag length to three as again indicated by the Akaike criterion and the final prediction error.

Overall the reactions remain similar to those from the full sample VAR model. However, figure 8 reveals a weaker reaction of the real exchange rate in the subsample. In the 2002–16 sample the responses of the unemployment rate and the Euro Stoxx 50 display the expected sign, although their responses are at no point significantly different from zero; see figure 5. In contrast to that, the sign of the reaction of the unemployment rate and the Euro Stoxx 50 in the 2008–16 sample is less clear, as the responses cross the zero line

Figure 8. Alternative Fifth Variable: Additional Real and Nominal Variables (2008:M10–2016:M10)

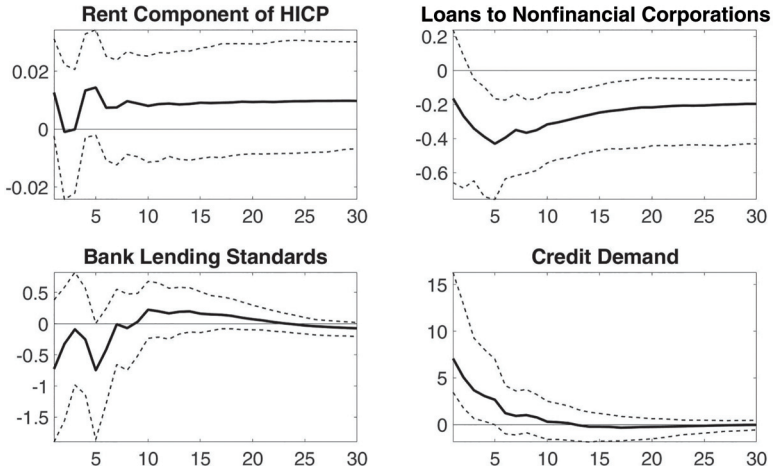


Notes: Responses of alternative choices for the fifth variable to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model estimated over the post-crisis sample with external instruments and 90 percent confidence band.

several times; see figure 8. Hence, we conclude that the policy transmission through employment and the stock market is particularly impaired in the post-2008 era. This era is characterized by sizable intra-euro-area government bond spreads, indicating that national characteristics play a major role for market participants during this time. This suggests that we can obtain more information from a country-specific perspective, which is pursued in the next section.

Figure 9 displays impulse responses for the credit market variables. Though not significant, the reaction of rent prices and lending standards are in line with the findings for the 2002–16 sample. In contrast, the increase in credit demand is substantially higher in the post-crisis sample. Interestingly, an expansionary monetary policy shock lowers the total loan volume to nonfinancial institutions. Our findings underpin the structural problems of the euro-area credit market: aggregate lending does not increase despite relaxed standards and higher credit demand.

Figure 9. Alternative Fifth Variable: Credit Market (2008:M10–2016:M10)



Notes: Responses of alternative choices for the fifth variable to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model estimated over the post-crisis sample with external instruments and 90 percent confidence band.

2.5 Pure Monetary Policy Shocks vs. Information Shocks

Jarocinski and Karadi (2018), Miranda-Agrippino and Ricco (2018), Nakamura and Steinsson (2018), and Romer and Romer (2000), among others, have pointed out that Governing Council decisions also unveil information about variables that do not represent policy instruments. The rationale behind this argument is that policymakers react to economic conditions (i.e., inflation and unemployment). An unanticipated decrease in the main refinancing rate might indicate that inflation is lower than expected and/or unemployment is higher than expected.

Jarocinski and Karadi (2018) disentangle the information component of monetary policy shocks from the pure policy shock by differentiating between responses of interest rates and stock prices on Governing Council meeting days. While interest rates decrease and stock prices increase after a monetary policy easing, both financial market variables move in tandem after an information shock.

Table 2. Loadings on the First Two Principal Components

Variable	PC 1	PC 2
German Government Bond 2Y	0.4433	−0.2401
German Government Bond 3Y	0.4699	−0.2176
German Government Bond 5Y	0.4744	−0.2082
German Government Bond 10Y	0.4272	−0.1683
Euro Stoxx 50	0.3028	0.6337
FTSE Euro 100	0.2887	0.6494

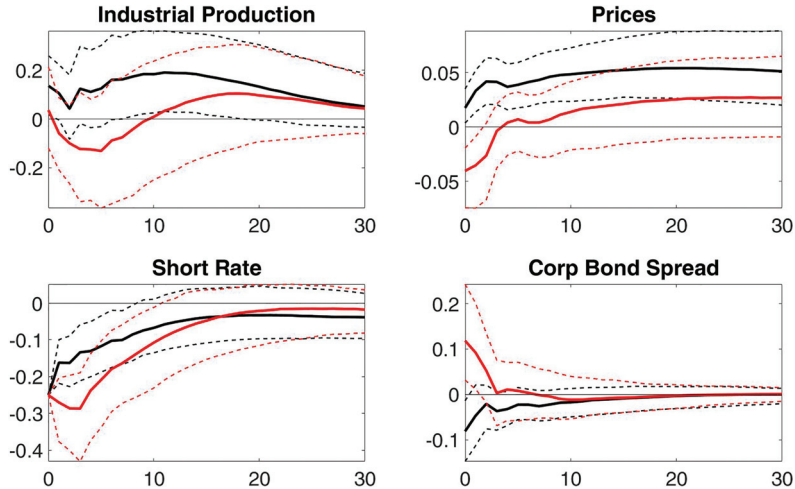
Therefore, Jarocinski and Karadi (2018) estimate a VAR model with instruments and impose sign restrictions in order to identify pure policy shocks and information shocks, respectively.

We follow their concept of discerning between movements of interest rates and stock prices on monetary policy announcement days. We incorporate changes in both variables into a principal component analysis. More precisely, we include standardized changes in the yield of German government bonds with maturities of 2, 3, 5, and 10 years on announcement days as well as standardized changes of the Euro Stoxx 50 and the Financial Times Stock Exchange (FTSE) Euro 100 stock price index.[16] By applying principal component analysis, we are more agnostic than Jarocinski and Karadi (2018), as we let the data speak without imposing restrictions. In fact, principal components and factor analyses are common empirical tools in the news announcement literature; see, e.g., Gürkaynak, Sack, and Swanson (2005) and Barakchian and Crowe (2013).

Table 2 displays the loadings on the first two components. The cumulative proportion of information explained by these two principal components is roughly 92 percent. While in the first component all variables are loaded with a positive sign, the second component loads changes in bond yields with a negative sign and changes in the stock market with a positive sign. Hence, we interpret the first component as the information shock and the second component as

[16]Each variable is standardized to have a mean of zero and a standard deviation of one.

Figure 10. Pure Policy and Information Shock



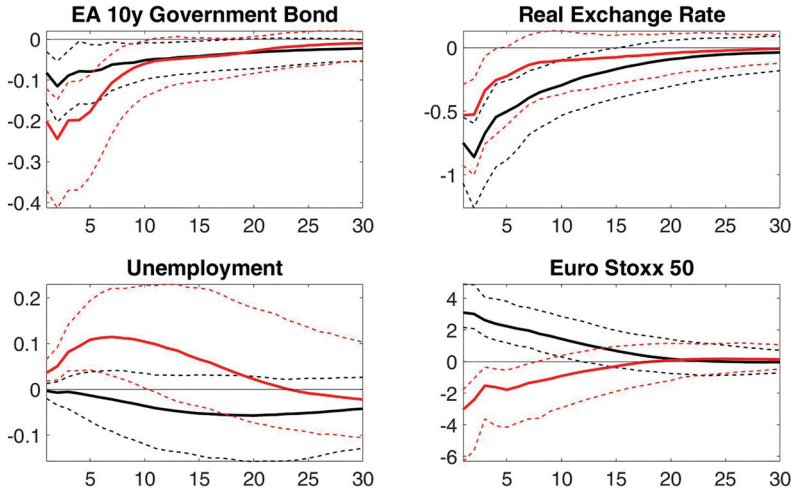
Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model estimated with external instruments and 90 percent confidence band. The black (red) line indicates the impulse responses for the pure monetary policy shock (information shock).

the pure policy shock.[17] The interpretation of the pure policy shock is further supported by the fact that the loadings on the lower end of the yield curve are higher (in absolute terms). According to the pure policy shock, the two biggest surprises were the decision not to raise the volume of the APP in December 2015 and the announcement of the SMP in May 2010.

Figure 10 shows the results from the baseline model following a pure policy shock and an information shock, respectively. The pure policy shock displays results which are qualitatively similar to those following the monetary policy shock discussed before. A negative information shock means that policymakers lower the interest rate due to weak economic fundamentals. In line with the findings of Jarocinski and Karadi (2018), a negative information shock decreases the price level and increases the corporate

[17] We multiply the pure policy shock with -1 such that, in line with the monetary surprise component from above, a positive surprise represents a monetary tightening.

Figure 11. Pure Policy and Information Shock: Additional Real and Nominal Variables

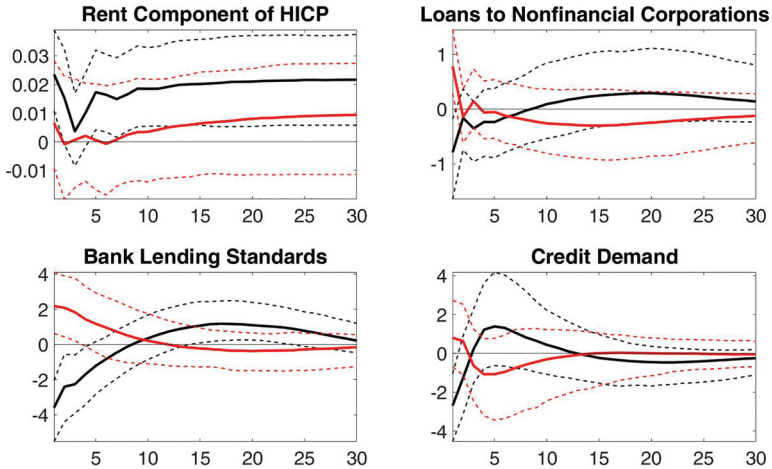


Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model estimated with external instruments and 90 percent confidence band. The black (red) line indicates the impulse responses to the pure monetary policy shock (information shock).

bond spread immediately. We do not find evidence that this shock decreases industrial production. Nevertheless, we observe a reduction in economic activity as indicated by the hike in unemployment; see figure 11. The stock market reacts in line with our expectations, i.e., a drop in the interest rate due to a pure policy shock (information shock) increases (decreases) the Euro Stoxx 50 on impact. A depreciation of the euro as well as decreases in government bond yields can be the consequence of both types of shocks.

From figure 12 we observe that rents only respond to a pure policy shock. As before, lending standards decrease and credit demand strengthens after a pure policy shock. A negative information shock leads to tighter lending standards but has no effect on credit demand. Once again, the loans to the private sector respond counter-intuitively with respect to both shocks. Apart from loans, all impulse responses to both types of shocks are well in line with the theory and the findings in the literature.

Figure 12. Pure Policy and Information Shock: Credit Market



Notes: Responses to an expansionary monetary policy shock of 25 bp obtained from the baseline VAR model estimated with external instruments and 90 percent confidence band. The black (red) line indicates the impulse responses to the pure monetary policy shock (information shock).

Overall, we find that the pure policy shock leads to impulse responses that are in line with those observed from the VAR model, where changes in the yield on 10-year German bunds on ECB meeting days serve as an instrument. Thus, we conclude that our instrument in the baseline model captures policy surprises well and is not overly distorted by the revelation of new information on meeting days.

3. Country-Specific Effects of Euro-Area Monetary Policy

In this section, we study the country-specific responses to a common monetary policy shock. Hence, at this stage, we want to exclude the feedback from domestic economic conditions to euro-area monetary policy. Since we have identified a common monetary policy shock in the previous section, there is no identification problem to solve at this stage. Therefore, in order to derive country-specific responses

to a common euro-area shock, we use local projections as suggested by Jordà (2005).

An impulse response is defined as the response of a variable h periods ahead to a monetary policy shock at time t. This response is not derived from a full-scale VAR model with interactions among all endogenous variables, but rather from a single-equation framework that does not allow for a feedback from the endogenous variable to monetary policy. We estimate a series of regressions of a dependent variable dated $t+h$ on the monetary policy shock in t as well as a set of control variables. The estimated model is the following:

$$y_{t+h} - y_{t-1} = \alpha_h + \beta_h mp_t^{EA} + \gamma_h' \sum_{s=0}^{q} \mathbf{x}_{t-s} + \varepsilon_{t+h}, \qquad (10)$$

where y_t is the dependent variable and $\mathbf{x}_t$ is a vector of country-specific control variables. We include up to q lags of control variables. The euro-area monetary policy shock is denoted by mp_t^{EA}. Hence, the coefficient β_h measures the impact of a change in policy at t on the dependent variable h periods ahead. Plotting β_h as a function of h provides us with an impulse response function.

For our purpose, local projections are advantageous for two reasons: (i) they rest on a very small number of parameters to be estimated and (ii) since we estimate a single equation only, the results are more robust to misspecifications in other parts of the model. While we typically model dynamic systems of equations, e.g., VAR models, *because* we want to capture the feedback from the economy to policy, we deliberately exclude this feedback here.

Due to the fact that the dependent variable is h periods ahead, the error terms will exhibit serial correlation. We therefore apply a Newey-West correction to our estimation errors, which we use to construct a confidence band around the estimated series of β_h coefficients. As suggested by Jordà (2005), the maximum lag for the Newey-West correction is set to $h+1$.

We estimate local projections for 10 member countries, which together account for more than 95 percent of euro-area GDP: Germany, France, Spain, Italy, Portugal, Greece, Ireland, Netherlands, Finland, and Austria. To contrast the country-specific responses with the area-wide responses, we also estimate the model for a

synthetic euro area that consists of these 10 countries only.[18] The sample period is 2002:M1 to 2016:M1 and the data frequency is monthly. The sample is slightly shorter than the sample used in the previous section due to limited data availability. We estimate the model for each of the following variables: (log) industrial production, (log) price level as measured by the HICP, unemployment rate, (log) real exchange rate, (log) stock prices, and (log) loans to the private sector.

We keep the list of control variables relatively short and include country-specific cyclical variables such as unemployment, prices, industrial production, and the real exchange rate. We also include the shadow short-term interest rate to reflect monetary conditions. Note that the latter is supposed to reflect the level of policy accommodation, but not the policy shock, which is reflected by mp_t^{EA}. Changing the vector of control variables has no substantive effect on our estimated impulse response functions.

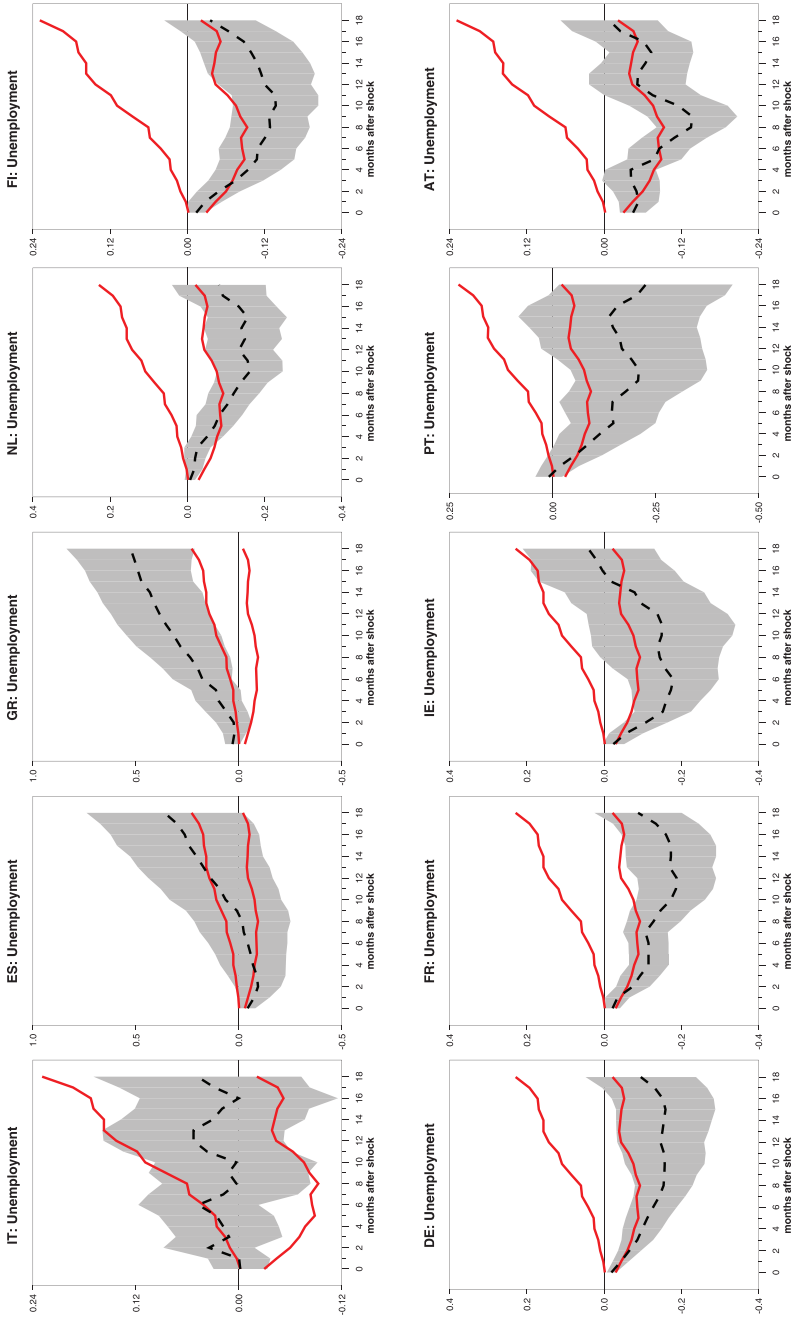
The euro-area monetary policy shock, mp_t^{EA}, is based on the identification of policy surprises discussed before. The relation between the structural shock u_t of the VAR model and the reduced-form shock ε_t is given by $U_t = S^{-1} \cdot \varepsilon_t$. From the estimation of the baseline model in 2.4, we receive the reduced-form error terms as well as the row in the matrix S^{-1} that captures the contemporaneous responses to the structural shock. With these variables at hand, we are thus able to uniquely identify our policy shock series, mp_t^{EA}.

3.1 Results

The results are presented in figures 13 to 18. In each figure, we plot the impulse response function following a monetary policy easing shock, the 90 percent error band around this impulse response, and, as a pair of red lines, the error band around the estimated impulse response for the synthetic euro-area variable. Thus, comparing the dotted country-specific impulse response and the red error

[18]The euro-area time series for each variable is constructed as the weighted average of the country-specific variables. For that purpose, the GDP weights from the ECB website have been normalized in order to account for member countries which are not included here, that is, the GDP weights for the 10 countries used here always add up to 100 percent.

Figure 13. Country-Specific Responses of Unemployment



Notes: Country-specific response to a euro-area monetary policy easing shock of 25 bp (dotted line) obtained from local projections and 90 percent error bands (shaded area). The solid lines are the error bands around the average euro-area response.

bands allows us to assess whether a given country's response deviates significantly from the response of the euro area as a whole.

We find that following a monetary policy shock, unemployment decreases significantly in core countries such as Germany, France, and the Netherlands; see figure 13. Some periphery countries—namely Italy, Spain, and Greece—in contrast, could not benefit from the monetary expansion implemented by the ECB. In these countries, unemployment does not fall. As a consequence of this heterogeneity, the area-wide unemployment rate does not respond significantly, which is consistent with the finding presented in the previous section.

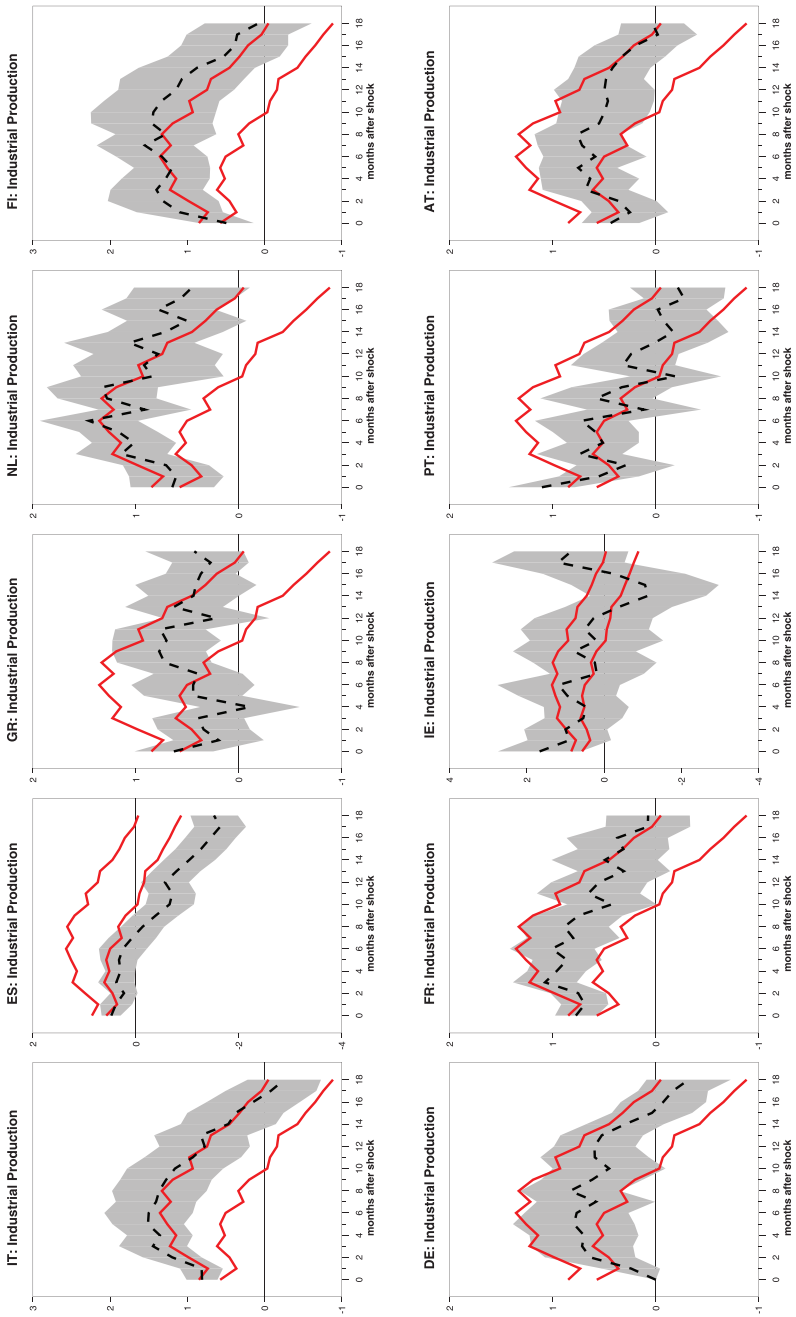
Figure 14 reports the responses of industrial production. Manufacturing activity improves in all countries following the expansionary policy shock. Interestingly, the responses are much more homogeneous across countries. Only the response of industrial production in Spain deviates markedly from the euro-area average response. In contrast to that, manufacturing activity rises in Italy and Greece after a monetary easing while employment does not improve. Possibly, unemployment rates in these countries are mainly driven by other sectors.[19] In a nutshell, we only observe heterogeneous effects on real activity if it is proxied by the unemployment rate and not by industrial production.

Figure 15 shows the responses of consumer prices, which increase moderately following a monetary policy shock. The responses are well in line with the average response of the euro-area price level and might be a result of the single European market.

In all countries, the real effective exchange rate depreciates on impact; see figure 16. Different responses of real exchange rates among EMU members can only occur when price levels react differently. As we find homogeneous responses of price levels to a monetary policy shock, we also observe homogeneous movements of the real exchange rate. The size of the depreciation for the euro area as a whole is in a range similar to the one observed in the VAR model (see figure 5). However, the confidence bands in the local projections framework are somewhat wider, suggesting that cutting the

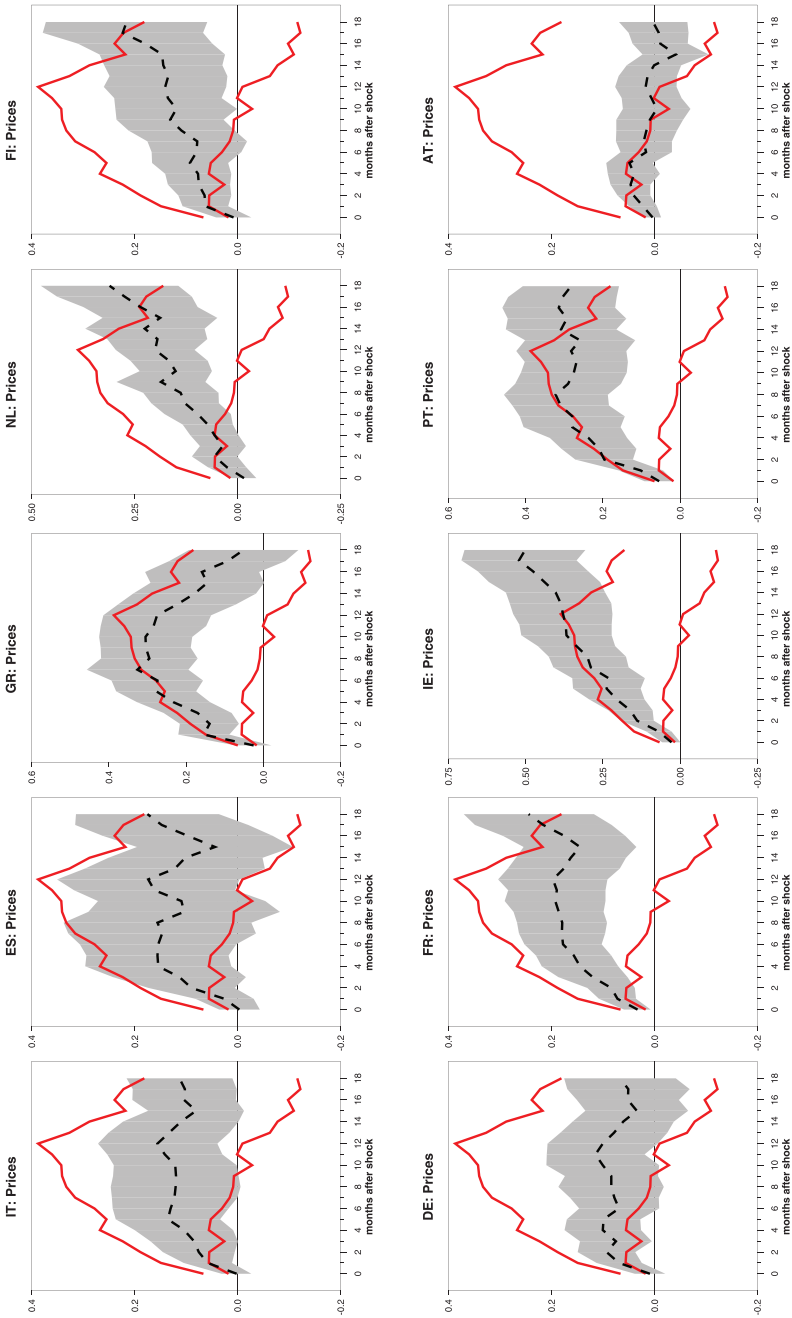
[19] In fact, the service sector accounts for more than two-thirds of GDP in both countries.

Figure 14. Country-Specific Responses of Industrial Production



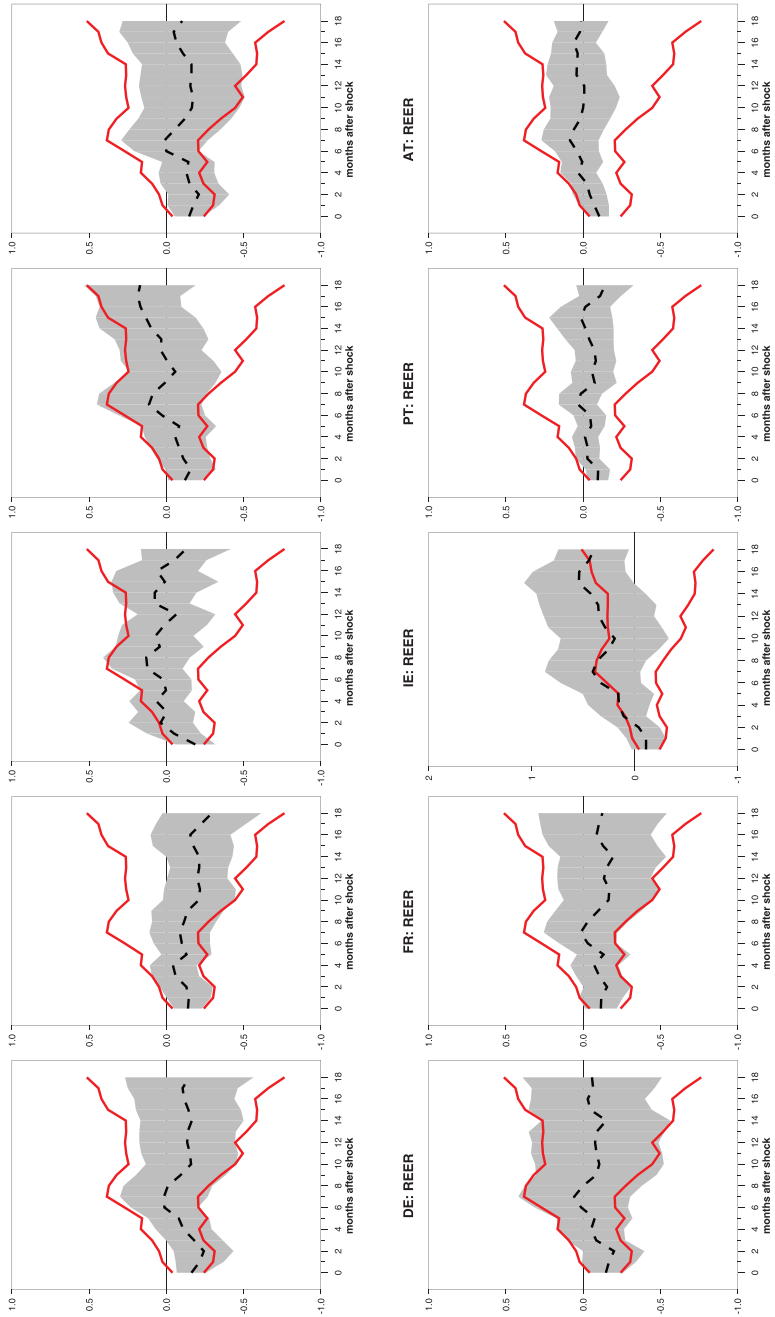
Notes: Country-specific response to a euro-area monetary policy easing shock of 25 bp (dotted line) obtained from local projections and 90 percent error bands (shaded area). The solid lines are the error bands around the average euro-area response.

Figure 15. Country-Specific Responses of Consumer Prices



Notes: Country-specific response to a euro-area monetary policy easing shock of 25 bp (dotted line) obtained from local projections and 90 percent error bands (shaded area). The solid lines are the error bands around the average euro-area response.

Figure 16. Country-Specific Responses of the Real Exchange Rate



Notes: Country-specific response to a euro-area monetary policy easing shock of 25 bp (dotted line) obtained from local projections and 90 percent error bands (shaded area). The solid lines are the error bands around the average euro-area response.

feedback from the real economy to monetary policy results in higher estimation uncertainty.

The response of the main stock price indexes (see figure 17) exhibits the expected positive sign for the overall euro zone in the short run. For some countries—i.e., Spain, Greece, the Netherlands, Ireland, Portugal, and Austria—the responses deviate negatively from the average euro-area response after about 10 months. The results are also in line with the country-specific findings provided by Wieladek and Pascual (2016). These authors also document an insignificant and even negative effect of a policy easing on stock prices in some euro-area countries.

Finally, figure 18 suggests that the ECB is not effective in stimulating credit to nonfinancial corporations. Germany, Austria, and Greece appear to be the only countries in which credit increases significantly following the monetary expansion. In most other countries, the response of bank credit remains insignificant. As derived from the VAR model and shown in figure 6, this is in line with the insignificant response of aggregate credit in the euro area.[20]

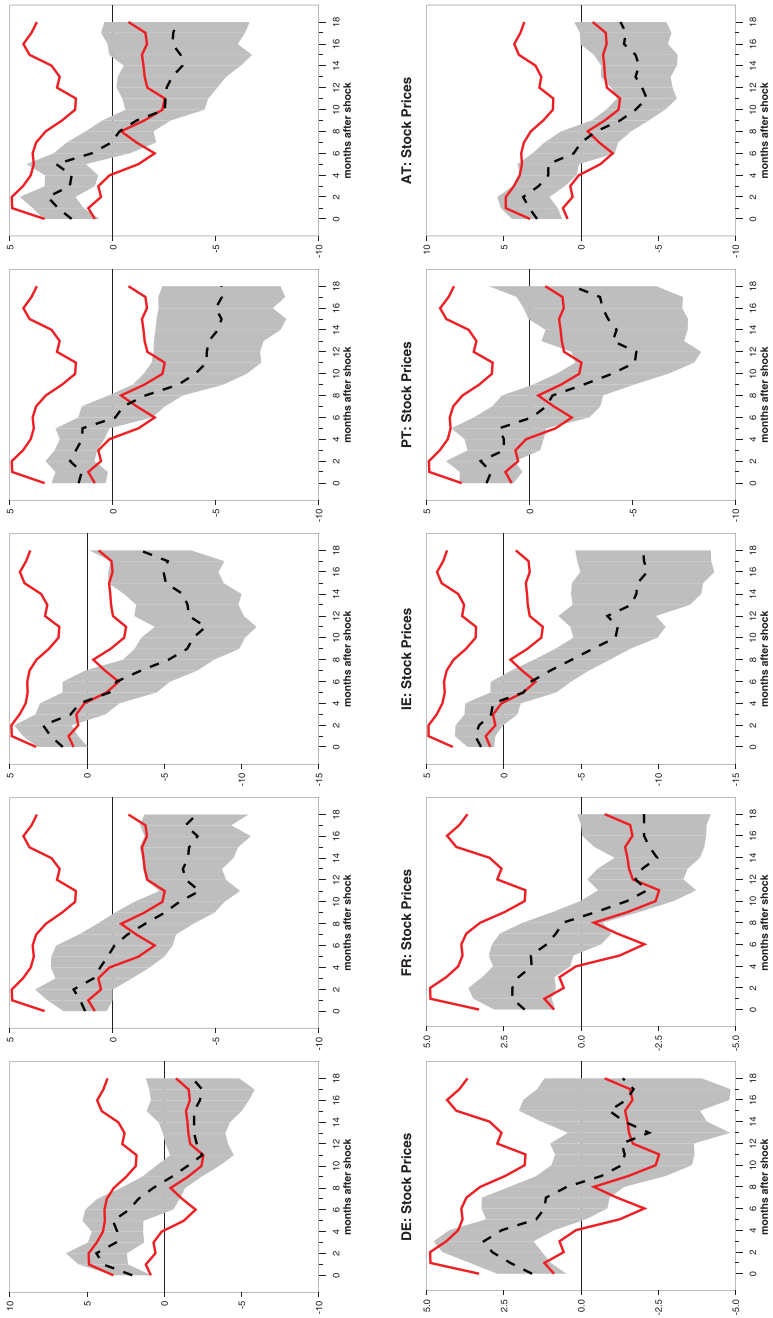
Overall, we find the responses of unemployment, stock prices, and bank lending to be different across member countries, while consumer prices and industrial production are much more homogeneous. The results suggest that the impaired transmission through the financial system, i.e., the stock market and the credit market, as well as structural frictions in the adjustment of the labor market, might hold the key to understanding the uneven transmission of ECB policy.

4. Conclusions

In this paper, we studied the monetary transmission mechanism in the euro area—both based on aggregate and country-specific data. To identify a monetary policy shock, we estimated an external instruments VAR that solves the contemporaneous correlation between monetary policy and financial variables in the euro area.

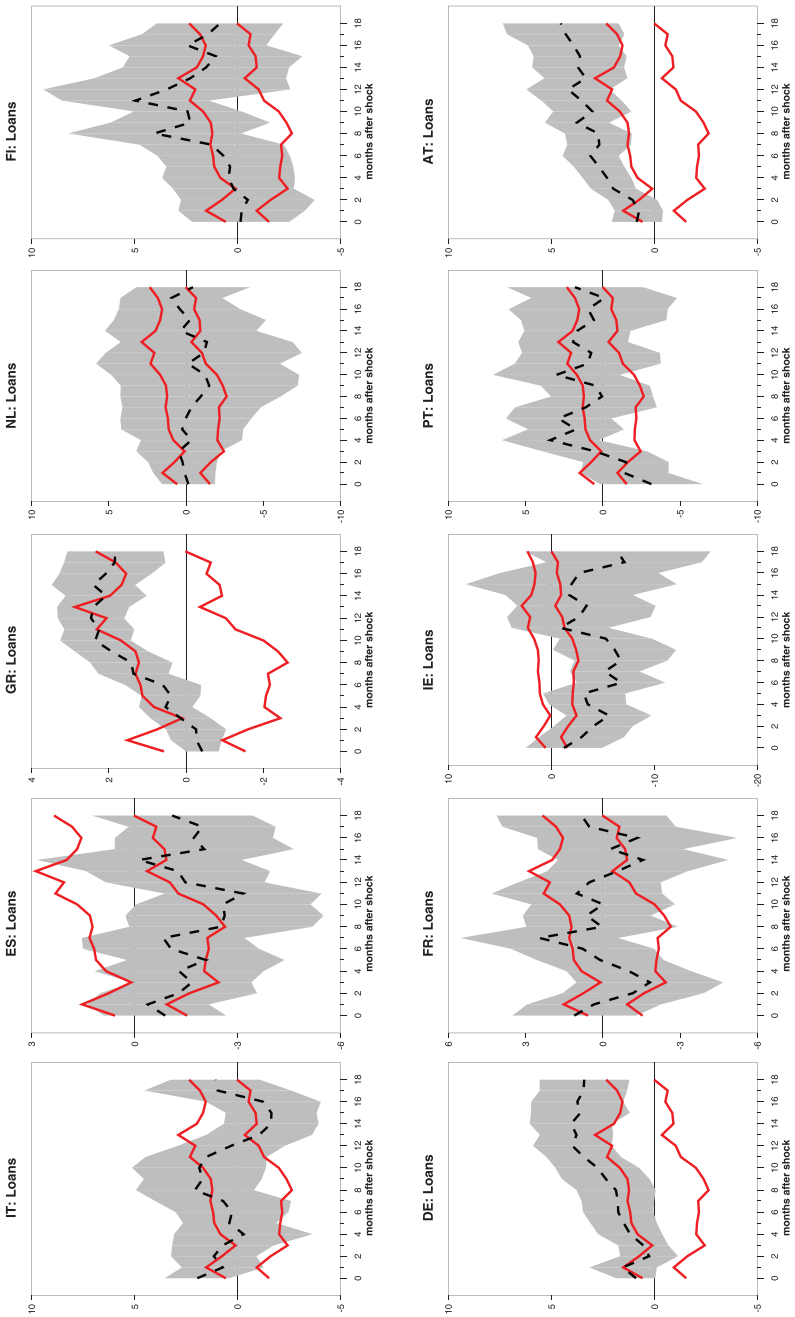
[20]This finding is in line with the results of Boeckx, Dossche, and Peersman (2017).

Figure 17. Country-Specific Responses of Stock Prices



Notes: Country-specific response to a euro-area monetary policy easing shock of 25 bp (dotted line) obtained from local projections and 90 percent error bands (shaded area). The solid lines are the error bands around the average euro-area response.

Figure 18. Country-Specific Responses of Loans



Notes: Country-specific response to a euro-area monetary policy easing shock of 25 bp (dotted line) obtained from local projections and 90 percent error bands (shaded area). The solid lines are the error bands around the average euro-area response.

Our findings are threefold: First, identifying a VAR with an external instrument helps to disentangle the simultaneous interaction of the ECB and the financial market. A principal component analysis of the responses of bond yields and stock prices combined with the interaction among the variables in the VAR model generates a plausible decomposition of policy surprises into the pure policy shock and the information shock arising from ECB decisions.

Second, we document the heterogeneity of the monetary transmission process across transmission channels. Overall, monetary policy is less effective with regard to stimulating bank lending and increasing the valuation of the stock market. These findings suggest that monetary transmission is severely hampered by the state of banking systems, e.g., the ongoing deleveraging and the burden of nonperforming loans.

Third, we shed light on the heterogeneity of policy transmission across member countries. For that purpose, we included the ECB's monetary policy shock in country-specific regressions. This makes sure that the policy shock is the same across countries and that a feedback from country-specific variables to euro-area monetary policy is excluded. We show that the responses of some variables, most notably prices and industrial production, are relatively similar across countries, while the transmission through the financial system, i.e., the responses of stock prices and bank lending, varies among member countries. Since our results are purely positive, we should be careful not to overemphasize the normative implications. Nevertheless, the results suggest that a "one-size-fits-all" monetary policy might not be the best tool to boost demand if national banking systems are blocked—not least since banks provide most financing in continental Europe. Over many years since the eruption of the European debt crisis, monetary policy was overburdened with the task of reviving economic activity. In light of the findings presented here, this has supported inflation throughout the euro zone. However, the effects on real activity are heterogeneous, especially if one focuses on unemployment, where core countries benefit from a monetary easing disproportionately.

Appendix. Data Sources and Definitions

Table A.1. Data Sources

Variable	Adj.	Source
Bank Lending Standards	nsa	Bank Lending Survey
Credit Demand	nsa	Bank Lending Survey
Crude Oil Prices (Brent Europe)	nsa	FRED
EONIA Rate	nsa	Datastream
EURIBOR Future	nsa	Eikon
Euro Stoxx 50	nsa	Datastream
Eurobond 10y All Ratings	nsa	ECB
FTSE Euro 100 Stock Price Index	nsa	Datastream
German Government Bond Yield	nsa	Eikon
Harmonised Index of Consumer Prices	sa	ECB
Industrial Production (excl. Construction)	sa	ECB
Loans to Nonfinancial Institutions	sa	ECB
Real Exchange Rate (vis-à-vis Group of 19 Trading Partners)	nsa	ECB
Shadow Rate	nsa	Wu and Xia (2016)
(Alternative) Shadow Rate	nsa	Krippner (2012)
(Alternative) Shadow Rate	nsa	Lemke and Vladu (2017)
S&P Eurozone Corporate Bond Yield	nsa	Datastream
Unemployment Rate	sa	ECB

Note: Seasonally adjusted data series are indicated by "sa." Data series not seasonally adjusted are indicated by "nsa."

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Contagion in the CoCos Market? A Case Study of Two Stress Events*

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The post-crisis regulatory framework has fostered the development of the market for contingent convertible bonds (CoCos). These instruments allow for loss absorption as a going concern, but their critics warn about their potential destabilizing effects in stress situations. We analyze the dynamics of the European CoCos market during two stress episodes that occurred in 2016 and were triggered by news on substantial unexpected losses faced by a European systemic bank. Our econometric approach aims at disentangling the fundamental contagion channels of the distress of such bank to the rest of the market from a potential CoCo-specific contagion channel. We find evidence of significant CoCo-specific contagion in the first stress episode that could result from investors' reassessment of CoCos' riskiness or from uncertainty on their supervisory treatment. We do not find instead evidence of CoCo-specific contagion in the second stress event, suggesting that as investors learn about the specificities of these instruments and their supervisory treatment, the CoCos market becomes more resilient.

JEL Codes: G14, G21, G28.

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1. Introduction

In the aftermath of the global financial crisis, bank capital requirements have been significantly increased to enhance the resilience of the system. In a number of jurisdictions the new regulatory framework allows banks to satisfy part of the additional capital needs with contingent convertible bonds (CoCos). The CoCos are hybrid debt instruments that in adverse contingencies can be used to recapitalize the banks by writing down their principal or converting them into equity.[1] Such a privileged regulatory treatment has spurred the growth of the CoCos market, with USD 521 billion of global issuances between the first issuance in 2009 and December 2015 (Avdjiev et al. 2020).

The rationale for the introduction of CoCos is related to their contingent debt-equity nature in two ways. First, they provide an equity injection when bank capital is needed and the bank is still a going concern, that is, when the bank is close to a given (regulatory) threshold but still above the regulatory minimums. Second, the debt feature, which prevails except in rare distress episodes, makes them appealing to both banks (as they carry a lower cost relative to equity) and fixed-income investors. Defendants of CoCos argue that these properties render these instruments optimal to increase the loss-absorption capacity of banks without imposing an excessive burden on bank profitability that could in turn negatively affect lending to the real economy (Duffie 2009; Squam Lake Working Group on Financial Regulation 2009; Flannery 2016; Vallee 2019). CoCos may also be beneficial by inducing timely recapitalization to avoid costly dilution of profits or loss of control (Pennacchi 2011; Calomiris and Herring 2013). However, these arguments are not shared by all academics and practitioners. Skeptics in fact argue that the double nature of CoCos introduces complexity and uncertainty that, should a CoCo be on the eve of conversion, may prove destabilizing for the issuer bank and could spill over to the rest of the market (Goodhart 2010; Allen 2012; Admati et al. 2013).

In this paper we shed light on the potential destabilizing effect of these securities by focusing on two stress periods experienced by

[1] Avdjiev, Kartasheva, and Bogdanova (2013) provide a description of the main design elements of CoCos.

Deutsche Bank (DB) during 2016 that created significant concerns about the value of its EUR 4.6 billion CoCo issuances. The two stress periods, which started in January and September 2016, respectively, and lasted for a few weeks each, were initially triggered by announcements of substantial legal costs to be faced by DB. During the two events, the yield to maturity (YTM) of DB's senior bonds and CoCos peaked to historical maximums. The unfolding of these events was accompanied by great concerns from practitioners and commentators about DB's CoCos, and in particular about both the possibility that their coupons would not be paid and that the supervisory authority could mandate their conversion. The CoCos issued by the most important European banks also experienced severe tensions, especially during the January episode when their average YTM attained historical maximums (figure 2A). In that event, the YTM of the senior bonds issued by the same European banks also increased, although the level remained relatively contained. These dynamics could be the result of the contagion from a systemically important institution such as DB to other banks through "fundamental channels" stemming from, e.g., the network of interbank exposures or the possibility of fire sales. Yet, the "overreaction" of the CoCo market relative to the behavior of the senior bond market suggests that, in addition to fundamental distress propagation, the CoCo market could have been further destabilized by a broad reassessment of the riskiness of these instruments, ignited by the specific problems of DB's CoCos. Such a CoCo-specific contagion is precisely the type of concern raised by skeptics on these instruments and was frequently emphasized by commentators of these events. Our analysis of the two DB stress periods attempts at identifying whether or not there has been a CoCo-specific contagion beyond any fundamental distress propagation. By doing so, we investigate the potential unintended financial stability effects of a regulatory regime that encourages banks to issue CoCos.

In our empirical analysis we regress the logarithm of the daily YTM of the CoCos of European banks (excluding DB) on (i) the institutions' stock return, the YTM of their senior bonds, and the daily change of their expected default frequency (EDF); and on (ii) the YTM of DB's CoCos and its interaction with a dummy for each of the two events. The first set of controls aims at capturing the extent to which the CoCos YTM can be explained by variables

related to the bank fundamentals such as its solvency risk (proxied by the YTM of the bank's senior bonds and by the EDF change) and its expected profitability (proxied by the bank's equity return). These variables allow us to control for the fundamental contagion from DB during the two distress episodes, and our specification thus captures how such contagion may have affected also CoCos prices. The second set of controls aims at capturing whether any interdependence between the Cocos of DB and those of other banks increased due to a Coco-specific contagion when DB was under distress.

Our results show across all the specifications used the existence of a CoCo-specific contagion channel in the propagation of the distress of DB to the rest of the European banks in the first of the two episodes, which supports CoCos skeptics' claim that these instruments might be a source of financial instability when negative shocks occur. Moreover, we also find that the behavior of the CoCos issued by riskier banks differs from those issued by safer institutions, which confirms also the existence of fundamental factors affecting the CoCo market dynamics. Yet, when comparing the CoCo-specific contagion between the two DB's events, we find that the channel is not at work anymore during the September episode. Such enhanced market stability could be consistent with investors' learning on the specificities of CoCos along time. The clarification by the European Banking Authority (2016b) and the European Central Bank (2016a) on the supervisory treatment of these instruments—in particular, with regard to the conditions that may call for the suspension of coupon payments—is also likely to have played a role in stabilizing the CoCos market during the second stress period of DB. All in all, the weakening of a CoCo-specific contagion mechanism suggests that the potential destabilizing role of these securities—claimed by CoCos skeptics—might be only transitory and/or could have been mitigated by regulators' intervention, but we are probably not yet able to say the last word.

The rest of the paper proceeds as follows. Section 2 briefly recalls the literature related to this paper. Section 3 describes the regulatory treatment of CoCos, the market development, and the two DB distress events. Section 4 discusses the two transmission channels of contagion: a fundamental and a CoCo-specific one. Section 5 describes the data. Section 6 presents the baseline empirical analysis. Section 7 provides some extensions of the baseline analysis. Section 8

presents robustness exercises. Section 9 discusses the policy implications of our findings and concludes.

2. Related Literature

Our paper belongs to a growing literature on CoCos, starting with Flannery (2005, 2016). The theoretical contributions initially dealt with how design features affect the valuation of these instruments (e.g., Posner 2010; Pennacchi 2011; Glasserman and Nouri 2012). The possibility of multiple pricing equilibriums for CoCos with market trigger and the ensuing manipulation incentives by the different stakeholders was uncovered in Sundaresan and Wang (2015), and subsequent contributions have identified design features that eliminate such pricing multiplicity (Calomiris and Herring 2013; Pennacchi, Vermaelen, and Wolff 2014). In addition to structural models, proposals for pricing CoCos are based on equity and credit derivatives (for a review, see Wilkens and Bethke 2014) and on Merton-type models (e.g., Brigo, Garcia, and Pede 2015). Another strand of the theoretical literature has focused on how wealth transfers between the bank's stakeholders upon CoCo conversion may affect the ex ante risk-taking decisions of the bank in a counterproductive way (Hilscher and Raviv 2014; Flannery 2016; Martynova and Perotti 2018). Closer to the focus of our study, concerns that CoCos may not offer effective loss absorption in times of distress have been raised by practitioners, academics, and regulators (Pazarbasioglu et al. 2011; Delivorias 2016).

The empirical literature on CoCos is scarce and has not addressed potential contagion problems associated with the conversion of these instruments. Berg and Kaserer (2015) investigate the impact of CoCos' issuances on banks' risk-taking using a sample of CoCo bonds issued by European banks over the period 2009–13. The paper finds that some key design features of CoCos, such as a too high conversion price, lead the issuer bank to increase risk-taking. The literature has also shown that riskier banks are less likely to issue CoCos (Goncharenko, Ongena, and Rauf 2020). In a similar vein, Avdjiev et al. (2020) show that the propensity to issue CoCos is higher for better capitalized banks.

This paper is related also to the literature on contagion. The theoretical and empirical debate on what contagion means is far from conclusive, as emphasized by Pericoli and Sbracia (2003) and Forbes

(2012). We interpret an increase in the co-movement of CoCo prices after controlling for variations in the banks' fundamentals as evidence of CoCo-specific contagion. The notion of contagion as a transmission of shocks in excess of what can be explained by fundamentals is discussed in Claessens, Dornbusch, and Park (2001), Forbes and Rigobon (2002), and Forbes (2012), and has been adopted in the analysis by, e.g., Calvo and Mendoza (2000), Bekaert, Harvey, and Ng (2005), and Bekaert et al. (2014).

3. The CoCos Market and the Two Deutsche Bank Stress Events

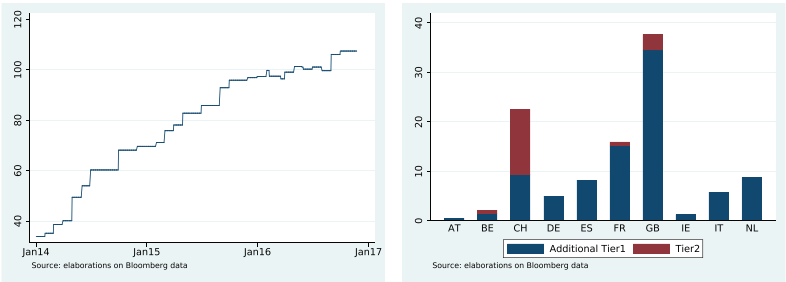
3.1 CoCos Market Developments and Regulatory Treatment

The issuance of CoCos by European banks has grown steadily since the adoption of Basel III and its introduction in Europe by means of the Capital Requirements Directive IV (CRD IV) and the Capital Requirements Regulation (CRR).

Under Basel III, hybrid capital instruments, such as CoCos, are designed to absorb losses while the bank is still a going concern and are classified as additional tier 1 (AT1) capital. To qualify as AT1, CoCos must have, like common equity, a perpetual maturity and noncumulative coupons, payable at the discretion of the issuer. They also need a loss-absorption trigger which allows the principal to be written down or converted to common equity without this event constituting a default of the issuer (Basel Committee for Banking Supervision 2011a). The trigger is activated when the common equity tier 1 (CET1) capital ratio of the bank falls below a CET1 ratio threshold that has to be set no lower than 5.125 percent (Basel Committee for Banking Supervision 2011b). These features should help to strengthen banks' capital position at a time when raising equity would otherwise be difficult.[2] Furthermore, allowing the issuer to miss coupon payments can reduce pressure on liquidity. Finally, in a resolution process, CoCos are senior only to common equity.

[2]In addition to the automatic triggers, supervisory authorities can mandate the conversion or the write-down of CoCos by discretionally activating the point of nonviability (PONV) triggers, if they believe that such action is necessary to prevent the bank from becoming insolvent (Delivorias 2016).

Figure 1. The Market of Contingent Convertible Bonds (EUR billion)



(a) European banks CoCos outstanding issuances (January 2014 - November 2016)

(b) European banks CoCos outstanding by country and Basel III capital category (November 2016)

In the European Union (EU), the CRD IV and the CRR make the coupon distribution on CoCos conditional on the banks' meeting the combined buffer requirement (CBR).[3] Should a bank not meet its CBR, the amount of payouts it can make in the form of dividends and coupon payments on AT1 instruments would be limited by the institution's maximum distributable amount (MDA) (CRD IV, Article 141).

The favorable regulatory treatment of CoCos—combined with banks' need to strengthen their capital position during difficult market conditions for equity issuances—has spurred the development of the CoCos market in Europe.[4] Based on the Bank of America Merrill Lynch Contingent Capital Index (BAMLCCI), European banks have almost tripled their CoCo issuances, with the outstanding volume rising from EUR 33.8 billion to EUR 107.4 billion between January 2014 and November 2016 (figure 1A).

The regulatory incentive in the growth of CoCos is confirmed by the fact that the vast majority of issuances qualify as AT1 capital, with only a small fraction of the total eligible as tier 2 capital

[3]According to Article 128 of the CRD IV, the CBR is the sum of the capital conservation buffer requirement and, when applicable, the countercyclical capital buffer requirement, the G-SII (global systemically important institutions) and O-SII (other systemically important institutions) buffer requirements, and the systemic risk buffer requirement.

[4]The need to strengthen capital adequacy has been felt by banks in particular during the preparation of the entry into force of the Single Supervisory Mechanism, which had as a milestone the completion of Asset Quality Review in October 2014.

(figure 1B). It is also worth highlighting that the tier 2 CoCos have mostly been issued in the early days of the newborn market well before the approval of Basel III.[5]

Despite this strong growth, toward the end of 2015, uncertainties emerged in the interpretation of the legal framework for the computation of the MDA trigger, i.e., the level of the capital requirements below which a bank would need to calculate the MDA and restrict its payouts, including those on AT1 instruments. In particular, it was unclear whether the CBR would come in addition to the 8 percent pillar 1 minimum requirements only, or in addition to both the pillar 1 and pillar 2 requirements. To clarify this issue, in December 2015 the European Banking Authority (2015) published an opinion in which it reaffirmed that both pillar 1 and pillar 2 requirements should be below the CBR when determining the MDA trigger. It was also recognized that, given the importance of the MDA trigger to investors in AT1 instruments, there should be full disclosure of the pillar 2 requirements. In early 2016, the European Central Bank (2016b) endorsed the EBA opinion. Notwithstanding such a clarification, uncertainties yet remained, this time due to the possibility that some banks could have breached the MDA trigger not because of actual losses but rather because of potential losses arising from the adverse scenario of the supervisory stress test, which would have been carried out over the following months.

Issuances of CoCos abruptly halted in early 2016, contemporaneously to the first DB stress period analyzed in this paper.

3.2 The Two Deutsche Bank Stress Events

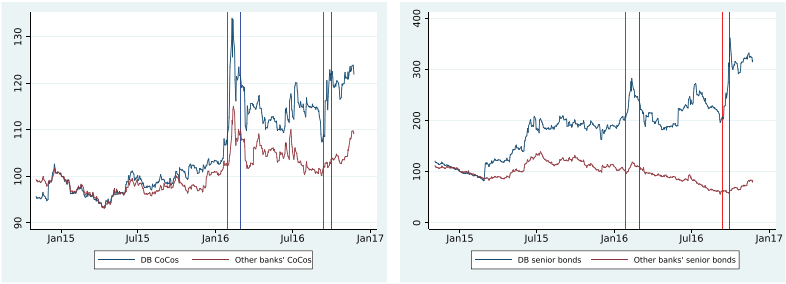
In this section we discuss in more detail each of the two distress episodes experienced by DB in 2016.

3.2.1 Event 1

The first event took place at the beginning of 2016 after the announcement of negative earnings expectations for 2015, which came as a surprise for market participants.

[5]The first CoCo was issued by Lloyds Banking Group in November 2009.

Figure 2. The CoCos and the Senior Bond Markets during Deutsche Bank's CoCos Distress (yield to maturity; January 2, 2015 = 100)



(a) The CoCos market (b) The senior bonds market

On January 20, 2016 DB announced that it expected an extraordinary high net loss of about EUR 6.7 billion for 2015, due to unanticipated very high litigation charges and restructuring costs and to cyclical market conditions (which led in particular to a fall in revenues from the bank securities trading unit). When the news was published, DB's debt securities reacted negatively (figures 2A and 2B). The situation worsened a few days later on January 28 when the chief executive officer (CEO) of Deutsche Bank, John Cryan, addressed a message to the bank's employees. In particular, he clarified that the yearly losses for 2015 (revised by then to EUR 6.8 billion) had been driven mainly by regulatory and litigation provisions (EUR 5.2 billion) and impairments on goodwill and other intangible assets (EUR 5.8 billion) while, differently from the previous announcement, little emphasis was given to the market conditions. The news about the loss, coupled with the uncertainty about the bank's capital position relative to the MDA trigger, raised dramatically the concerns on the capability of DB to meet the forthcoming payments on its CoCos (Glover 2016). This in turn led the bank to take a number of initiatives to restore market confidence: on February 8, 2016, in a press release and in a message to the bank employees by the chief financial officer (CFO) of Deutsche Bank, Marcus Schenck, the bank's availability of cash to make the upcoming payments due on the CoCos was reaffirmed.[6] Then on the

[6]http://www.db.com/newsroom_news/2016/ir/deutsche-bank-publishes-updated-information-about-at1-payment-capacity-en-11391.htm and

following day another message to the employees was sent by the CEO, who claimed that "Deutsche Bank remains absolutely rock-solid."[7] On February 12, 2016, the CFO announced a tender offer to buy back in the market some of the bank senior unsecured debt, "taking advantage of market conditions to repurchase this debt, lowering its debt burden at attractive prices" and with "no impact on the bank's capacity to service coupons on its AT1 capital."[8] The buyback operation was clearly aimed at proving the soundness of the bank to the markets. On February 29, 2016, the CFO announced the success of the operation, which was welcomed by the market.[9] After that date, conditions began to improve. The stabilization of the CoCos market might have also benefited from the recognition by the ECB, in February 2016, of the need to reduce uncertainty on the supervisory treatment of these instruments, and, in particular, on the conditions that would lead to the suspension of coupon payments. After long debate, the ECB announced on July 29, 2016, changes in the methodology to calculate the banks' MDA trigger that would effectively reduce the likelihood that banks would not be allowed to make CoCo coupon payments (see section 9 for more details).

Before the described event, the price of DB's CoCos was essentially stable. However, at the beginning of 2016 the YTM started to rise with a sudden spike of almost 200 basis points (figure 2A). In particular, the YTM jumped from an average of around 660–670 basis points in mid-January 2016 to more than 850 basis points in mid-February. We set the start date of the first distress episode on January 28 when, after the profitability announcement and the first message by the CEO of DB, the YTM of DB's CoCos spiked by more than 100 basis points. We set the end date on February 29

http://www.db.com/newsroom_news/2016/ghp/a-message-from-marcus-schenck-deutsche-bank-s-additional-tier-1-at1-capital-en-11386.htm.

[7]http://www.db.com/newsroom_news/2016/ghp/a-message-from-john-cryan-to-deutsche-bank-employees-0902-en-11392.htm.

[8]http://www.db.com/newsroom_news/2016/ghp/a-message-from-marcus-schenck-deutsche-bank-announces-a-public-tender-offer-to-buy-back-debt-en-11399.htm.

[9]http://www.db.com/newsroom_news/2016/medien/deutsche-bank-announces-results-of-public-tender-offer-for-us-dollar-bonds-as-of-early-tender-date-en-11494.htm.

when, after the announcement of the successful completion of a debt buyback, the YTM substantially declined.[10]

3.2.2 Event 2

The second distress event of DB started on September 12, 2016 when a number of articles in the press announced that the bank had received a USD 14 billion fine by the U.S. Justice Department. The news came after a series of other bad news that involved the bank over the summer,[11] putting the institution under serious investors' scrutiny that was further exacerbated by the reports that the German government wouldn't help the ailing bank (Donahue 2016). Both DB's CoCos and senior bonds yields took a hit, with the senior bonds exceeding the peak observed in February 2016 (figures 2A and 2B).

The rise in DB senior bonds and CoCos YTM started therefore on September 12, with an average increase on all issues by about 10 basis points in one day for both types of securities. Then other large one-day increases followed, particularly on September 16 when the news of the fine was confirmed—with the average YTM of DB senior bonds and CoCos increasing by 36 and 49 basis points, respectively (Davies 2016 and Strasburg 2016)—and on September 26 when the rumors about the government unwillingness to help the bank spread (Donahue 2016)—with the average YTM of DB senior bonds and CoCos increasing by 22 and 29 basis points, respectively. A number of interventions at the end of September contributed to the recovery from this distress. In particular, on September 30 the CEO of DB reassured the market and the bank employees about the soundness

[10]The results presented in section 6.3, section 7, and section 8 are robust to setting the beginning of the distress event to January 20, i.e., when DB announced for the first time the negative results.

[11]The most relevant are the following: (i) in June the International Monetary Fund (IMF) stated that DB "appears to be the most important net contributor to systemic risks in the global banking system" (IMF 2016); (ii) also in June, the U.S. subsidiary of DB failed for the second consecutive year the stress test performed by the Federal Reserve, as the regulator rejected the bank capital plan (Federal Reserve Board 2016); (iii) in July, the rating agency Standard & Poor's lowered its outlook on the bank's rating; (iv) again in July, the bank passed the EU-wide stress test by a tight margin (EBA 2016a); and (v) in August the bank was removed from the Stoxx 50 index.

of the bank's financial conditions;[12] on the same day, Eurogroup President Jeroen Dijsselbloem stated that the USD 14 billion fine to DB was excessive,[13] and unconfirmed reports claimed that the bank was about to agree on a reduction of the fine with the U.S. Department of Justice.[14] Taking this information into account, in our empirical analysis we set September 12 as the start date of the second distress event, and September 30 as the end date.

4. The Transmission Channels: Fundamental and CoCo-Specific Contagion

The sequence of news in January and September 2016 led to mounting concerns on the ability of DB to pay the coupons on its CoCos or even of a possible write-down of their notional value. This information also increased the insolvency risk of DB, as shown by the large increases in the YTM of its senior debt. During both periods stock prices declined significantly, reflecting mounting investors' concerns on DB's profitability outlook. As exhibited in figures 2A and 2B, the rest of the European banking sector experienced also some distress: the YTM of the senior bonds and of the CoCos of European banks increased (and their stock prices fell). Such a comovement may be the result of a propagation mechanism based on two contagion channels: a *fundamental* one and a *CoCo-specific* one.

First, DB is a global systemically important institution whose financial difficulties may spill over to other institutions. This could result from the losses that DB's counterparties would suffer directly in case of its insolvency, and indirectly from the exposure to other institutions through a loss cascade propagated by the network of interbank exposures. The distress of a systemic intermediary with important trading activities such as DB may also lead to substantial fire sales that would depress asset values and negatively affect the capitalization of banks with similar assets in their trading portfolio.

[12]http://www.db.com/newsroom_news/2016/medien/media-reports-on-deutsche-bank-a-message-to-employees-from-john-cryan-en-11702.htm.

[13]http://euobserver.com/tickers/135321.

[14]The rumors were indeed confirmed by the facts, as an agreement had been found in December when the fine was almost halved to USD 7.2 billion.

Negative spillovers on the economy could also arise due to second-round general equilibrium effects following the failure of a systemic institution. For example, it could reduce or make more costly the access to credit for banks' customers, which may as a response cut down investment or hiring; this in turn could depress the rest of the economy and, through this channel, the quality of other banks' assets. We refer to all these propagation mechanisms as "fundamental." They may explain all or part of the price evolution of banks' liabilities, including their CoCo instruments.

Second, both in January and September 2016, the news that DB may face substantial losses immediately brought the attention to the institutions' capability to satisfy the coupon payments of its CoCos and to the possibility that supervisors would prohibit the bank from paying their coupons. More generally these events led to statements by practitioners and commentators alike on a broad reassessment of the riskiness of CoCos, and to concerns that their design and their debtlike features during normal times could have led investors to misunderstand these instruments. During these stress episodes, investors' uncertainty about CoCos could have led to adverse price dynamics extending from DB's CoCos to the CoCos issued by other institutions, well beyond what would be explained by fundamental factors (e.g., banks' credit risk, profitability, etc.) We refer to such a propagation as "CoCo-specific contagion."

The European banking sector was affected by the difficulties of DB in the two stress episodes, albeit to a different extent. While the severity of the distress of DB, as proxied by the change in the YTM of its senior bonds, was of a similar intensity in the two periods, the magnitude of the movements in the other European banks' senior bonds and CoCos YTM was significantly smaller in the second event.

In the first event, the yield of the senior bonds of the major European banks experienced a short-lived increase, within the range of variation observed in the previous 12 months. This suggests a—possibly weak—fundamental contagion from DB to the other major European banks. In contrast, the European banks' CoCos yield hit an all-time high, which constitutes suggestive evidence of a CoCo-specific contagion.

In the second event, instead, the senior bond yield of the European banks remained essentially unchanged, which we can interpret as evidence of a very weak or even lack of fundamental contagion.

The CoCos yields increased somewhat, although their maximum level remained well below that attained in the first event. The lower magnitude of the movements in the CoCos market during the second stress episode relative to the first could be the result of a lower fundamental contagion and/or a lower CoCo-specific contagion. A less intense contagion through the Coco-specific channel could be due to an ameliorated understanding by investors of the mechanisms and risks underlying these instruments following the January episode. Furthermore, the actions taken by the ECB during the summer, namely the reduction of the capital ratio below which coupon payments would not be allowed by regulators, could have contributed as well to mitigate volatility transmission through this channel.

In the next sections we investigate whether the described anecdotal evidence on both a fundamental and a CoCo-specific contagion, as well as their relative magnitude across the two events, can be confirmed by the empirical analysis.

5. Data

The main data set used is provided by the BAMLCCI. This index tracks on a daily basis the performance of contingent capital bonds publicly issued in the major markets. In addition to closing market data (e.g., price, YTM, spread, and duration), information from this source includes a large number of variables at International Securities Identification Number (ISIN) level (e.g., maturity date, rating, amount outstanding, issuance currency, and whether the instrument is classified as AT1 or T2 capital). Further information on CoCos' main contractual characteristics is extracted from Bloomberg, including for instance the loss-absorption mechanisms (principal write-down/equity conversion), the type of write-down (temporary, permanent), the underlying trigger variable (e.g., tier 1 ratio, CET1), and the trigger level. Data at issuer level are collected from various sources: (i) senior unsecured bonds are from the Merrill Lynch index ER00; (ii) daily (closing) equity prices are from Bloomberg; (iii) five years expected default frequency (EDF) probability are from Moody's Analytics, and (iv) balance sheet data are

Table 1. Summary Statistics of CoCos Issuances

	No. of Daily Observations	No. of Issues	No. of Issuers	Volume Outstanding (EUR Billion, Nov. 2016)
Austria	105	1	1	0.5
Belgium	1,276	2	1	2.1
France	7,036	15	3	15.9
Germany	2,930	5	2	5.0
Italy	2,848	6	2	4.3
The Netherlands	1,070	3	2	3.1
Spain	4,189	9	3	10.5
Switzerland	8,168	16	2	21.7
United Kingdom	13,902	29	4	32.3
Total	41,524	86	20	95.4

from S&P Global Market Intelligence.[15] The securities in our sample are issued by European banks between June 2014 and November 2016.[16] Overall we have a sample of 86 CoCos issued by 20 banks located in nine European countries (table 1).[17] The U.K. and the Swiss banking systems are the most represented in our sample, with both the largest number of issuances (29 and 16 issues, respectively) and the highest volume outstanding (about EUR 32 billion and EUR 22 billion, respectively, as of November 2016). Securities in the data set are denominated in EUR, USD, and GBP. Tables 2, 3, and 4

[15]Matching the original data set with stock prices, EDF, and balance sheet information has led to a reduction of the original panel of banks from the BAMLCCI index, which originally included data on 98 CoCos and 28 issuers located in 10 countries.

[16]The empirical specification includes the daily average YTM of the CoCos issued by DB (see section 6). Regressions therefore start in June 2014 when a CoCo issued by DB enters the BAMLCCI.

[17]Information on issuers is aggregated at group level if the holding bank is represented in our sample. The banks included are the following: Erste Group (Austria), KBC Bank (Belgium), Societe Generale, Credit Agricole, and BNP Paribas (France), Aareal Bank and Deutsche Bank (Germany), Intesa SanPaolo and Unicredit (Italy), ABN Amro and Ing Group (the Netherlands), Banco Bilbao Vizacaya Argentaria, Bankinter, and Banco Santander (Spain), Credit Suisse and UBS (Switzerland), Barclays Plc., HSBC, Lloyds Banking Group, and Royal Bank of Scotland (United Kingdom).

Table 2. Summary Statistics of CoCos Issuances by Loss-Absorption Mechanism

	Number of Issues				Volume Outstanding (EUR Billion, November 2016)			
	Temporary Write-Down	**Permanent Write-Down**	**Equity Conversion**	**Total**	**Temporary Write-Down**	**Permanent Write-Down**	**Equity Conversion**	**Total**
Austria	1	—	—	1	0.5	—	—	0.5
Belgium	1	1	—	2	1.4	0.7	—	2.1
France	14	1	—	15	15.2	0.7	—	15.9
Germany	5	—	—	5	5.0	—	—	5.0
Italy	6	—	—	6	4.3	—	—	4.3
The Netherlands	1	—	2	3	1.0	—	2.1	3.1
Spain	—	1	8	9	—	1.0	9.4	10.5
Switzerland	—	15	1	16	—	21.7	—	21.7
United Kingdom	—	2	27	29	—	3.1	29.2	32.3
Total	28	20	38	86	27.4	27.3	40.8	95.4

Table 3. Summary Statistics of CoCos Issuances by Trigger Variable

	Number of Issues				Volume Outstanding (EUR Billion, November 2016)			
	CET1 Ratio	**Core Tier 1 Ratio**	**Total Risk-Based Capital Ratio**	**Total**	**CET1 Ratio**	**Core Tier 1 Ratio**	**Total Risk-Based Capital Ratio**	**Total**
Austria	1	—	—	1	0.5	—	—	0.5
Belgium	2	—	—	2	2.1	—	—	2.1
France	15	—	—	15	15.9	—	—	15.9
Germany	5	—	—	5	5.0	—	—	5.0
Italy	4	—	2	6	4.0	—	0.2	4.3
The Netherlands	3	—	—	3	3.1	—	—	3.1
Spain	9	—	—	9	10.5	—	—	10.5
Switzerland	16	—	—	16	21.7	—	—	21.7
United Kingdom	23	6	—	29	32.3	—	—	32.3
Total	78.0	6.0	2.0	86.0	95.2	0.2		95.4

Table 4. Summary Statistics of CoCos Issuances by Trigger Level

	Number of Issues				
	Low Trigger			High Trigger	
	5%	5.125%	6%	7%	Total
Austria	—	1	—	—	1
Belgium	—	1	—	1	2
France	—	14	—	1	15
Germany	—	4	—	1	5
Italy	—	4	2	—	6
The Netherlands	—	1	—	2	3
Spain	—	8	—	1	9
Switzerland	7	4	—	5	16
United Kingdom	6	—	—	23	29
Total	13	37	2	34	86

Notes: We follow Avdjiev, Kartasheva, and Bogdanova (2013) and define a high (low) trigger CoCo bond as one having a trigger level above (below) 6 percent of the ratio between common equity tier 1 capital and risk-weighted assets (CET1/RWA).

present detailed summary statistics regarding the main contractual features for the selected sample of CoCos. Statistics by country of the average daily YTM of the CoCos and of the senior bonds as well as of the average returns of equity prices, of average EDF change, and of distance to trigger are provided in table 5.

6. Empirical Analysis

6.1 Setup

In order to disentangle fundamental from CoCo-specific contagion, we define the following baseline model:

$$\begin{aligned}\log(Y_{i,j,t}) = \alpha_{i,j} + \beta \log(Y_{i,j,t-1}) + \sum_{k=1,3} \gamma_k X_{k,j,t} + \eta C_{i,j,t} + \delta Y_{DB,t} \\ + \sum_{s=1,2} \varrho_s Dummy_{s,t} \\ + \sum_{s=1,2} \varphi_s Y_{DB,t} Dummy_{s,t} + \theta Z_{i,j,t} + \varepsilon_{i,j,t}. \qquad (1)\end{aligned}$$

Table 5. Additional Statistics on CoCos, Senior Bonds, Stocks, EDF, and Distance to Trigger

	CoCo Bonds (Avg. YTM, bp)			Senior Bonds (Avg. YTM, bp)			Stocks (Avg. Returns, bp)			EDF 5Y (Avg. Change, bp)			Distance to Trigger (Avg., bp)		
	2014	**2015**	**2016**	**2014**	**2015**	**2016**	**2014**	**2015**	**2016**	**2014**	**2015**	**2016**	**2014**	**2015**	**2016**
Austria	—	—	854	—	—	137	—	—	26	—	—	−0.61	—	—	810
Belgium	601	595	601	61	104	105	4	9	1	0.09	−0.01	0.03	742	974	886
France	676	665	685	155	135	92	−8	4	2	−0.06	−0.13	−0.15	506	530	600
Germany	615	652	742	168	179	219	−13	−4	−13	0.04	−0.01	−0.06	929	841	724
Italy	710	711	771	162	117	94	−10	2	−30	0.21	−0.12	0.37	797	734	720
Netherlands	—	626	610	—	108	87	—	−6	4	—	−0.01	0.03	—	631	770
Spain	700	693	746	105	113	82	−8	−12	−3	0.02	0.02	0.02	624	678	682
Switzerland	524	535	550	90	62	61	−7	2	−12	−0.04	0.02	0.08	1,111	1,147	1,085
United Kingdom	640	609	662	169	145	114	−1	−3	−1	−0.11	−0.04	0.10	556	660	581

The log-linear specification of the model helps to control for the possible nonlinearity that exists in the relations between senior bonds and CoCos, and between equity prices and CoCos, which can be especially important when banks' financial position is eroded and the pricing of their CoCos becomes more information sensitive. The dependent variable, $\log(Y_{i,j,t})$, is the natural logarithm of the daily YTM of CoCo i, issued by a European bank j other than DB, at time t. We include as a regressor one lag of the dependent variable, $\log(Y_{i,j,t-1})$, to account for serial correlation in yields. The bank-specific controls $X_{k,j,t}$, with $k = 1, 2, 3$, consist of the daily average YTM of the senior bonds ($k = 1$) issued by bank j, of the daily stock return ($k = 2$) of bank j, and of the daily change in the bank EDF five years ahead, all at time t. By using these controls, we aim to capture the variation in the CoCos YTM that can be explained by the changes in a bank's fundamental conditions such as its solvency (proxied by the YTM of the bank's senior bonds and the bank EDF) and its expected profitability (proxied by the bank's equity return). To the extent that there is a fundamental contagion from DB to the other banks' CoCos YTM during the two distress episodes, it should be reflected by these variables. The CoCo and bank-specific variable $C_{i,j,t}$ is the distance to trigger, namely the difference between the bank j capital ratio and the CoCo i trigger level, measured at time t. As a measure of a CoCo's risk, we expect it to matter for its pricing. This variable is updated quarterly (the regulatory reporting frequency for bank capital) and the information becomes available with a lag to the reference date.

The variable $Y_{DB,t}$ is the daily average YTM of the CoCos issued by DB which we use to control for the possible co-movement between the CoCos of DB and those of the other banks.[18] In particular, the coefficient of $Y_{DB,t}$ in the baseline regression would capture the average interdependence in the YTM of CoCos, along all the sample and time span considered, which could result from unobserved variables that affect the pricing of all CoCos, but not that of the other bank securities.

We interact $Y_{DB,t}$ with the dummy variables $Dummy_{s,t}$, with $s = 1, 2$, which identify the two DB stress periods as previously

[18]The results in the paper are robust to an alternative specification with the natural logarithm of the YTM of DB's CoCos instead of the YTM of DB's CoCos.

defined. The coefficients of the interactions aim to capture whether there has been a change in the interdependence between the CoCo of DB and those of other banks when DB was under distress (and investors spotted the riskiness of its CoCos).

Finally, Z_i is a vector of CoCo-specific variables that includes the type of action upon conversion (principal write-down or equity conversion), the type of write-down (temporary or permanent), the regulatory treatment of the instrument (AT1 or tier 2), the trigger level (high or low[19]), the possibility for the supervisory authority to suspend coupon payments, and the size of the issue.

We use the baseline regression to test the following hypothesis:

Hypothesis 1. *Existence of CoCo-specific contagion in the first and/or the second stress period.*

We say that there is evidence of CoCo-specific contagion in the stress period s if the coefficient of $Y_{DB,t}Dummy_{s,t}$ is statistically significant and positive. Our definition of contagion is consistent with that of Forbes and Rigobon (2002) that considers contagion to be an increase in cross-dependence across markets during crises periods. Note that since we control for the banks' equity returns, senior unsecured debt YTM, and EDF change, the potential fundamental contagion from DB to the other banks' CoCos should be transmitted through the effect of DB's distress on these variables.

6.2 Estimation

Equation (1) is first estimated using a generalized least squares (GLS) estimator with random effects, to control for the issue-specific characteristics of the CoCos. We estimate the model by clustering standard errors by a "bank∗time" variable and by allowing for intragroup correlation (i.e., relaxing the usual assumption that the observations are independent), since the null hypothesis of no serial correlation in the residuals is rejected by the Wald test discussed

[19]We follow Avdjiev, Kartasheva, and Bogdanova (2013) and define a high (low) trigger CoCo bond as one having a trigger level above (below) 6 percent of the ratio between common equity tier 1 capital and risk-weighted assets (CET1/RWA). The results of the analysis do not change if we set the threshold at 5.125 percent, the regulatory minimum for a CoCo to qualify as AT1.

in Wooldridge (2002) and for which Drukker (2003) shows the good size and power properties.

In addition, as the likelihood-ratio test shows that there is heteroskedasticity across groups, we also run a feasible GLS (FGLS) estimation with heteroskedastic error structure across panels and a panel-specific AR(1) process within panels.

The time dimension of our panel is large enough (638 days) and hence we a priori expect the dynamic panel bias of GLS and FGLS estimations to be small. We nonetheless check the results against the generalized method of moments estimator proposed by Blundell and Bond (1998), also known as system GMM. The latter framework accounts for endogeneity, controls for unobserved heterogeneity, and deals with the biases and inconsistencies typical of least-square estimations, provided that the model is not subject to serial correlation of order two and the instruments used are valid. We estimate the system GMM with a two-step estimator robust to panel-specific autocorrelation and heteroskedasticity.

6.3 Results

The results from the estimation of equation (1) are shown in table 6, with the GLS estimation reported in column 1, the FGLS in column 2, and the system GMM in column 3. They show, as expected, that the CoCo YTM is inversely related with the bank stock returns, as shown by the significance of the stock return variable (columns 1 and 2), and that it positively co-moves with the senior bonds YTM (columns 1 and 2) and with the EDF change (columns 1 and 2). The negative sign of the coefficient of the distance to trigger (columns 1 and 2) indicates as well that the CoCos issued by the riskier banks, i.e., banks with a relatively lower capital ratio, are priced at a higher YTM. These findings show that there is a close relationship between the banks' fundamental conditions and the pricing of their CoCos. Any effect of the two DB events on the other banks' CoCos resulting from the fundamental contagion channel previously described should therefore be captured by the significance of the senior bond, stock return, and EDF variables. This is less the case for the distance to trigger, which—although being a significant driver of CoCos YTM—has a much lower frequency (quarterly instead of daily as in the case of the other variables) and is observed with a long lag.

Table 6. Fundamental and CoCo-Specific Dependence of CoCos YTM during DB Events

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Lag CoCo YTM	0.994***	0.998***	0.993***
	(0.002)	(0.000)	(0.002)
Senior Bond YTM	0.166**	0.039***	0.052
	(0.075)	(0.012)	(0.175)
Stock Return	−0.094***	−0.092***	−0.077
	(0.012)	(0.002)	(0.091)
EDF Change	1.223***	0.402***	6.138
	(0.319)	(0.118)	(5.123)
Distance to Trigger	−0.024**	−0.009***	−0.013
	(0.012)	(0.003)	(0.010)
CoCo DB	0.039	0.052***	−0.299
	(0.029)	(0.008)	(0.225)
CoCo DB × Event1	1.227***	1.414***	4.085**
	(0.205)	(0.048)	(1.809)
CoCo DB × Event2	−0.050	−0.093	−0.026
	(0.124)	(0.082)	(0.110)
Event1	−0.095***	−0.108***	−0.315**
	(0.015)	(0.004)	(0.139)
Event2	0.005	0.007	0.004
	(0.009)	(0.006)	(0.008)
CoCos Characteristics:			
Subordination Type	−0.004*	−0.001***	−0.005***
	(0.002)	(0.000)	(0.001)
Equity Conversion	−0.000	0.000	−0.001
	(0.001)	(0.000)	(0.001)
Permanent Write-Down	0.002*	0.000**	0.002
	(0.001)	(0.000)	(0.001)
Cancellation by Regulators	−0.001*	−0.000**	−0.001**
	(0.001)	(0.000)	(0.001)
Trigger Level	0.001	0.000**	−0.001
	(0.001)	(0.000)	(0.001)
Size	0.000	−0.000	0.001
	(0.001)	(0.000)	(0.001)
Constant	−0.020***	−0.010***	0.002
	(0.005)	(0.001)	(0.017)

(*continued*)

Table 6. (Continued)

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Observations	37,987	37,987	37,987
Number of ISIN	82	82	82
Hansen Test			80.07
p-value			0.114
Arellano-Bond Test for AR(1)			−1.838
p-value			0.066
Arellano-Bond Test for AR(2)			0.611
p-value			0.541
Number of Instruments			83

Notes: Estimates of the yield to maturity of the CoCos issued by European banks, excluding DB. Random-effect GLS estimation with clustered standard errors by "bank*time" and intragroup correlation is shown in column 1; feasible GLS estimation with heteroskedastic error structure across panels and panel-specific AR(1) process within panels is reported in column 2; Blundell-Bond system GMM estimate is shown in column 3. For the GMM, Hansen, AR(1), and AR(2) tests are provided. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

Table 6 shows also that there is limited evidence of co-movement between the CoCos of DB and those of the other banks over the entire time span considered, as the DB's CoCos variable is significant only under the FGLS estimation (column 2). Such significance could be due to unobservable variables that affect all CoCos and that might bias the estimates. In section 8.2 we therefore analyze the robustness of the FGLS results with respect to the possibility of omitted variables.

As for the variables of direct interest to test the CoCo-specific contagion hypothesis described in the previous paragraph, we find that the interaction $Y_{DB,t}Dummy_{s,t}$ is positive and very significant across all three alternative estimations during the first stress period (i.e., $s = 1$); the coefficient of the interacted variable is not different from zero in the second episode (i.e., $s = 2$), which suggests that no CoCo-specific contagion took place in the second event. This confirms the hypothesis about the existence of CoCo-specific contagion during the first stress episode only.

The results for the CoCo-specific contagion during the first distress event are also economically significant, albeit to a different extent in the three estimates. The coefficients of the interaction term imply that a 100 basis points higher YTM of the CoCos of DB during the first distress event translates into an average increase of the YTM of the CoCos issued by the other European banks between 1.23 percent (according to the GLS estimate) and 4.09 percent (according to the GMM estimate). With simple algebra it can be shown that the CoCo-specific contagion from DB to the other banks has explained between 18 and 61 percent of the observed increase in the average CoCo YTM of the other banks in the first stress event (i.e., between January 28, 2016 and the peak of the event on February 11, 2016).

The estimated CoCo-specific contagion is not only significant but could be also a lower bound. It is in fact identified as the increase of CoCo yields observed when DB was under distress, additional to what is explained by the fundamental contagion. The assumption does not however consider any negative impact that CoCos yields may have on bank equity prices, one of the variables used to proxy fundamental contagion. This feedback is, however, at least theoretically possible according to Brigo, Garcia, and Pede (2015).

As for the CoCo-specific characteristics, there is evidence of their role in explaining the YTM of CoCos, with four out of five variables being significant under at least one of the estimations. The issue-specific subordination (AT1 versus tier 2) is the more relevant CoCo-specific variable, as it is significant across all three estimates, with the negative sign of the coefficient indicating that the more senior the CoCos, the lower their YTM (the higher its price). In terms of loss-absorption mechanism, the variable capturing the alternatives of equity conversion versus principal write-down is not significant, while the variable controlling for the type of principal write-down (i.e., permanent versus temporary) is significant (columns 1 and 2), with the expected positive sign indicating that the CoCos facing the possibility of a permanent writing down have on average a higher YTM. The possibility for the regulator to cancel the interest payments also matters, as the corresponding variable is significant across all the estimations. The negative sign of the coefficient indicates that the CoCos that allow for supervisory judgment in deciding about the suspension of interest payments have a lower YTM (higher price).

This suggests that a certain degree of flexibility in the decision to impose losses on bondholders is considered to be a risk-reduction factor by investors. Supervisors might have in fact some tolerance in deciding to impose losses if, for instance, the bank has a credible plan to strengthen its solvency. The trigger level is also significant, but only column 2 and its positive coefficient confirm the intuition that the CoCos with a higher trigger have a higher YTM than the CoCos with a lower trigger. The size of the issuance has instead no role in explaining the YTM (and the price) of the CoCos.

Finally, the value of the coefficient of the lag of the CoCo YTM—approximately one—raises some concerns about the presence of a unit root in the CoCos YTM process that could bias our results. We test for this possibility with the Fisher-type test, in essence a Phillips-Perron test robust to general forms of heteroskedasticity (Maddala and Wu 1999).[20] The test rejects the null hypothesis of all panels having a unit root at 1 percent.[21] This notwithstanding, to get reassurance about the robustness of the results in table 6 to the presence of a unit root for some CoCos, we reestimate equation (1) on the subsample of CoCos for which the test excludes the possibility of unit root (by doing so we include CoCos from 11 out of 20 banks). The results are broadly unchanged.[22]

7. Further Analysis of CoCo-Specific Contagion

In the previous section we have found evidence on the existence of CoCo-specific contagion during the first DB distress episode that

[20]The characteristics of our data make testing for nonstationarity nontrivial: (i) heteroskedasticity poses important limitations to the reliability of test statistics for several tests; (ii) the time horizon considered is somewhat shorter than what is typically required for unit-root testing; (iii) the unbalanced structure with gaps of our panel limits the available tests.

[21]In performing the test we face the usual tradeoff related to the number of lags to be used, between bias (with low number of lags) and variance/errors (with increasing number of lags). In our case the choice should go to zero or a low number of lags, as in financial markets, under the efficient market hypothesis, it should not be possible to use past information to forecast future yields. In any case, this tradeoff does not seem to play an important role for our data, as with lags from 0 to 10 the test always rejects the null hypothesis of all panels having a unit root against the alternative of at least one panel being stationary.

[22]The results of the Fisher-type test and for this robustness test are not shown for brevity but are available on request.

is not explained by the dynamics of banks' fundamental risk. In this section, we further analyze the determinants and implications of contagion.

7.1 Bank Riskiness and CoCos Dynamics

To further explore the interaction between fundamental and CoCo-specific contagion, we augment the baseline model in equation (1) by the following three triple interactions: (i) $Y_{DB,t}Senior_{j,t}\ Dummy_{s,t}$, (ii) $Y_{DB,t}EDF_{j,t}\ Dummy_{s,t}$, and (iii) $Y_{DB,t}C_{i,j,t}\ Dummy_{s,t}$.

A positive coefficient in the first two interactions and/or a negative coefficient for the third interaction would imply a higher sensitivity of the price of the CoCos issued by the riskier banks to the developments of DB. Such type of behavior would be the one to expect if the dynamics of CoCos were driven by fundamentals, as the CoCos issued by riskier banks are more likely to suffer restrictions on coupon payments or a conversion. In contrast, nonsignificant coefficients would indicate that the interdependence between European banks' CoCos and those issued by DB does not depend on the riskiness of the issuer and would reinforce the evidence of CoCo-specific contagion. The results of the analysis are presented in table 7. The significance of the interactions of the EDF (columns 1 and 2) and of the distance to trigger (column 2) provides suggestive evidence about a connection between the CoCos market dynamics and the risk of the issuing banks, and points toward the possible existence of a fundamental contagion in the CoCos dynamics in addition to the CoCo-specific contagion during the first DB stress period, which is confirmed by the interaction $Y_{DB,t}Dummy_{s,t}$. In addition, during the second DB stress event, when the Coco-specific contagion seems to be dispelled—as already shown in table 6—there is still a fundamental contagion at work in the CoCos dynamics, as confirmed by the significance of the interaction of the EDF across all estimations.

7.2 Banks' Location, Business Model, and CoCo-Specific Contagion

To further shed light on the sources of CoCo-specific contagion in the first event, we consider the role played by banks' business model and banks' location. In particular, we investigate whether (i) banks

Table 7. Fundamental and CoCo-Specific Dependence of CoCos YTM during DB Events: Bank Risk and CoCos YTM Robustness

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Lag CoCo YTM	0.994***	0.998***	0.993***
	(0.002)	(0.000)	(0.002)
Senior Bond YTM	0.166**	0.039***	0.024
	(0.076)	(0.012)	(0.189)
Stock Return	−0.086***	−0.089***	−0.116
	(0.012)	(0.002)	(0.107)
EDF Change	0.600*	0.115	0.910
	(0.348)	(0.121)	(6.224)
Distance to Trigger	−0.018	−0.007***	−0.018
	(0.012)	(0.003)	(0.013)
CoCo DB	0.036	0.050***	−0.348
	(0.029)	(0.008)	(0.264)
CoCo DB × Event1	1.280***	1.389***	3.836***
	(0.171)	(0.050)	(1.213)
CoCo DB × Event2	−0.056	−0.111	0.032
	(0.128)	(0.083)	(0.281)
Event1	−0.090***	−0.104***	−0.311***
	(0.015)	(0.004)	(0.105)
Event2	0.006	0.008	−0.002
	(0.009)	(0.006)	(0.019)
Senior Bond × Event1 × CoCo DB	−1.592	−0.129	2.710
	(2.166)	(0.513)	(4.445)
Senior Bond × Event2 × CoCo DB	1.445	1.395	2.521
	(1.767)	(1.300)	(4.406)
EDF Change × Event1 × CoCo DB	114.151***	64.578***	−51.154
	(21.025)	(4.581)	(46.834)
EDF Change × Event2 × CoCo DB	78.747***	35.436**	685.167***
	(29.293)	(15.705)	(231.847)
Distance to Trigger × Event1 × CoCo DB	−1.298	−0.370***	2.204
	(1.167)	(0.113)	(1.977)
Distance to Trigger × Event2 × CoCo DB	−0.303	−0.069	−0.667
	(0.310)	(0.180)	(0.509)
CoCos Characteristics:			
Subordination Type	−0.004*	−0.001***	−0.005***
	(0.002)	(0.000)	(0.002)
Equity Conversion	−0.000	0.000	−0.001
	(0.001)	(0.000)	(0.001)
Permanent Write-Down	0.002*	0.000**	0.001
	(0.001)	(0.000)	(0.001)

(*continued*)

Table 7. (Continued)

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Cancellation by Regulators	-0.001^{*}	-0.000^{**}	-0.002^{**}
	(0.001)	(0.000)	(0.001)
Trigger Level	0.001	0.000^{**}	-0.001
	(0.001)	(0.000)	(0.001)
Size	0.000	-0.000	0.001
	(0.000)	(0.000)	(0.001)
Constant	-0.020^{***}	-0.010^{***}	0.007
	(0.005)	(0.001)	(0.020)
Observations	37,987	37,987	37,987
Number of ISIN	82	82	82
Hansen Test			77.49
p-value			0.308
Arellano-Bond Test for AR(1)			-1.823
p-value			0.068
Arellano-Bond Test for AR(2)			0.358
p-value			0.720
Number of Instruments			95

Notes: Estimates of the yield to maturity of the CoCos issued by European banks, excluding DB. Random-effect GLS estimation with clustered standard errors by "bank*time" and intragroup correlation is shown in column 1; feasible GLS estimation with heteroskedastic error structure across panels and panel-specific AR(1) process within panels is reported in column 2; Blundell-Bond system GMM estimate is shown in column 3. For the GMM, Hansen, AR(1), and AR(2) tests are provided. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

located in the same region as DB, i.e., within the euro area, or (ii) banks with a business model closer to that of DB, i.e. investment banking, might have experienced higher contagion. As for the first test, we define a "euro-area" (EA) dummy, which is equal to 1 if a CoCo has been issued by a bank located in the euro area, and zero otherwise. We then augment equation (1) with EA_j and with its interaction with $Y_{DB,t}Dummy_{s,t}$. The results in table 8 indicate that the YTM of the CoCos issued by euro-area banks are on average higher than those issued by non-euro-area banks, as shown by the positive and significant coefficient of the EA_j dummy in the three estimates (columns 1–3). However, the CoCos issued by the euro-area banks do not behave differently vis-à-vis those

Table 8. Fundamental and CoCo-Specific Dependence of CoCos YTM during DB Events: The Role of Bank Location

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Lag CoCo YTM	0.993***	0.997***	0.992***
	(0.003)	(0.000)	(0.002)
Senior Bond YTM	0.177**	0.047***	−0.016
	(0.078)	(0.012)	(0.182)
Stock Return	−0.094***	−0.090***	−0.121
	(0.012)	(0.002)	(0.094)
EDF Change	1.208***	0.449***	5.048
	(0.324)	(0.117)	(4.011)
Distance to Trigger	−0.016*	−0.009***	−0.001
	(0.009)	(0.003)	(0.015)
CoCo DB	0.037	0.058***	−0.345
	(0.030)	(0.009)	(0.251)
CoCo DB × Event1	1.213***	1.457***	4.822**
	(0.222)	(0.049)	(2.000)
CoCo DB × Event2	−0.052	−0.070	0.173
	(0.124)	(0.083)	(0.197)
Event1	−0.095***	−0.111***	−0.369**
	(0.015)	(0.004)	(0.152)
Event2	0.005	0.006	−0.011
	(0.009)	(0.006)	(0.014)
Euro-Area Bank	0.001*	0.001***	0.002**
	(0.001)	(0.000)	(0.001)
Euro-Area Bank × Event1 × CoCo DB	0.031	−0.008	−0.105
	(0.046)	(0.006)	(0.073)
Euro-Area Bank × Event2 × CoCo DB	0.009	0.001	0.018
	(0.011)	(0.008)	(0.021)
CoCos Characteristics:			
Subordination Type	−0.003*	−0.001***	−0.005***
	(0.002)	(0.000)	(0.001)
Equity Conversion	0.000	0.000*	−0.000
	(0.000)	(0.000)	(0.001)
Permanent Write-Down	0.003*	0.001***	0.002
	(0.001)	(0.000)	(0.001)
Cancellation by Regulators	−0.001	−0.000**	−0.001**
	(0.000)	(0.000)	(0.001)
Trigger Level	0.000	0.000	−0.002
	(0.001)	(0.000)	(0.001)
Size	0.001	−0.000	0.001
	(0.001)	(0.000)	(0.001)
Constant	−0.023***	−0.013***	0.003
	(0.007)	(0.001)	(0.018)

(*continued*)

Table 8. (Continued)

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Observations	37,987	37,987	37,987
Number of ISIN	82	82	82
Hansen Test			78.95
p-value			0.099
Arellano-Bond Test for AR(1)			−1.834
p-value			0.0666
Arellano-Bond Test for AR(2)			0.796
p-value			0.426
Number of Instruments			84

Notes: Estimates of the yield to maturity of the CoCos issued by European banks, excluding DB. Random-effect GLS estimation with clustered standard errors by "bank*time" and intragroup correlation is shown in column 1; feasible GLS estimation with heteroskedastic error structure across panels and panel-specific AR(1) process within panels is reported in column 2; Blundell-Bond system GMM estimate is shown in column 3. For the GMM, Hansen, AR(1), and AR(2) tests are provided. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

issued by the non-euro-area banks during the DB stress periods, as shown by the nonsignificance of the triple interaction. Hence the bank's location has no role in the transmission of the CoCo-specific contagion.

As for the second test, using information from banks' balance sheet at the end of 2016, we consider a bank to be an "investment bank" (IB) if its trading assets account for more than 30 percent of its total assets. We then define a dummy variable IB_j which captures whether a bank j's business model is IB or not, and estimate again equation (1) augmented by IB_j and its interaction with $Y_{DB,t}Dummy_{s,t}$. The results in table 9 show that investment banks might face on average a higher CoCo YTM, as indicated by the positive coefficient of the IB_j dummy in column 3. More importantly in the context of the analysis, they also indicate that investment banks experience more severe CoCo-specific contagion during the first DB distress event, as shown by the significance of the coefficient of the triple interaction $IB_jY_{DB,t}Dummy_{s,t}$ (column 2).

Table 9. Fundamental and CoCo-Specific Dependence of CoCos YTM during DB Events: The Role of Bank Business Model

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Lag CoCo YTM	0.994***	0.997***	0.987***
	(0.002)	(0.000)	(0.005)
Senior Bond YTM	0.162**	0.042***	0.224
	(0.076)	(0.012)	(0.203)
Stock Return	−0.094***	−0.092***	−0.104*
	(0.012)	(0.002)	(0.062)
EDF Change	1.192***	0.395***	4.370
	(0.329)	(0.118)	(3.076)
Distance to Trigger	−0.020*	−0.008***	0.103
	(0.011)	(0.003)	(0.064)
CoCo DB	0.036	0.053***	−0.318
	(0.029)	(0.008)	(0.238)
CoCo DB × Event1	1.221***	1.411***	4.454**
	(0.209)	(0.048)	(1.854)
CoCo DB × Event2	−0.047	−0.094	0.254
	(0.123)	(0.082)	(0.227)
Event1	−0.095***	−0.108***	−0.343**
	(0.015)	(0.004)	(0.142)
Event2	0.004	0.007	−0.017
	(0.009)	(0.006)	(0.016)
Investment Bank	0.001	0.000	0.024**
	(0.000)	(0.000)	(0.009)
Investment Bank × Event1 × CoCo DB	0.047	0.018***	−0.047
	(0.037)	(0.007)	(0.047)
Investment Bank × Event2 × CoCo DB	0.003	0.001	0.039
	(0.012)	(0.009)	(0.030)
CoCos Characteristics:			
Subordination Type	−0.004*	−0.001***	−0.012**
	(0.002)	(0.000)	(0.006)
Equity Conversion	−0.000	0.000	0.004
	(0.001)	(0.000)	(0.004)
Permanent Write-Down	0.002*	0.001**	0.008
	(0.001)	(0.000)	(0.005)
Cancellation by Regulators	−0.001*	−0.000**	−0.007*
	(0.001)	(0.000)	(0.004)
Trigger Level	0.001	0.000**	0.002
	(0.001)	(0.000)	(0.003)
Size	0.000	−0.000	−0.000
	(0.001)	(0.000)	(0.003)
Constant	−0.020***	−0.011***	−0.028
	(0.006)	(0.001)	(0.020)

(*continued*)

Table 9. (Continued)

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Observations	37,987	37,987	37,987
Number of ISIN	82	82	82
Hansen Test			78.67
p-value			0.249
Arellano-Bond Test for AR(1)			−1.826
p-value			0.068
Arellano-Bond Test for AR(2)			0.669
p-value			0.504
Number of Instruments			91

Notes: Estimates of the yield to maturity of the CoCos issued by European banks, excluding DB. Random-effect GLS estimation with clustered standard errors by "bank*time" and intragroup correlation is shown in column 1; feasible GLS estimation with heteroskedastic error structure across panels and panel-specific AR(1) process within panels is reported in column 2; Blundell-Bond system GMM estimate is shown in column 3. For the GMM, Hansen, AR(1), and AR(2) tests are provided. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

7.3 Heterogeneity in CoCo-Specific Contagion

The results discussed so far describe the average features of our sample of CoCos issued by European banks. An interesting question is how "representative" these numbers are: does the CoCo-specific contagion during the DB events differ significantly and systematically across CoCo issues? The large-T nature of our panel allows us to answer this question by estimating issue-specific model and scrutinizing the cross-sectional distributions of the coefficients. The results confirm the findings of the panel estimation. Figure 3 shows the distribution of the coefficients of $Y_{DB,t}Dummy_{s,t}$ obtained from ordinary least squares (OLS) estimates of equation (1) performed ISIN by ISIN. The estimated coefficients appear to be a good summary of the data, as the banks' CoCo YTM are indeed positively correlated with the YTM of the DB CoCos in the first event. Coherently with the panel estimates, the distribution of the estimated coefficients for the first turbulent period has both mean and median at about 1.3, while that for the second period has both a negative mean (at about −0.6) and a negative median (at about −0.4), with a negative skew.

Figure 3. Cross-Sectional Distribution of the Estimated CoCo × DB Interacted Coefficients

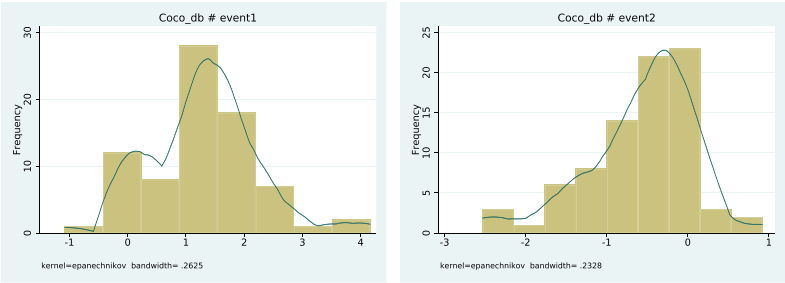


Table 10. Significance of CoCo-Specific Contagion across Issue-Specific OLS Regressions

Estimated Significant Coefficients (Share of Total Estimates)		
Significance Level	**CoCo DB × Event1**	**CoCo DB × Event2**
Up to 1 Percent	0.54	0.14
1 to 5 Percent	0.27	0.13
Up to 5 Percent	0.80	0.26

The share of significant coefficients of $Y_{DB,t}Dummy_{s,t}$ across the OLS estimates of equation (1) performed ISIN by ISIN is shown in table 10. The share of significant coefficients (at either the 1 or 5 percent level) of the ISIN-by-ISIN estimates is much higher in the first stress event, with 80 percent of the estimated coefficient being significant. The statistically significant coefficients involve almost all issuers (84 percent). The estimated coefficients are instead significant, with a much lower frequency in the second stress event (26 percent of the times) and for less than half of the banks in the sample (42 percent). What is also worth noting is that the coefficients relative to the second event, when significant, have always a negative sign, while in the first event all but two of the higher number of estimated coefficients have a positive sign. The key message is therefore that our panel estimations provide a good description for most CoCo issues in the sample.

7.4 Financial Stability Implications of CoCo-Specific Contagion

In this section we study the effects of the distress of DB and of CoCo-specific contagion on the insolvency risk of the banks in the sample. This allows us to assess the financial stability implications of the CoCo-specific contagion. Using the results in section 7.3, we construct a new bank variable, Exp_j, which is the average at bank level of the coefficients of $Y_{DB,t}Dummy_{s,t}$ estimated across all the CoCos i issued by bank j. The average is computed under the convention that nonsignificant or negative coefficients are set to zero and the significantly positive coefficients are set at their estimated value. A positive value of Exp_j indicates that a bank is exposed to the first DB distress via the CoCo-specific contagion channel, and the higher the value, the higher the exposure. We then estimate the following equation:

$$\begin{aligned} Y_{i,j,t} = \alpha_{i,j} &+ \beta Y_{i,j,t-1} + \delta log(Y_{DB,t}) + \gamma Dummy_{1,t} \\ &+ \eta log(Y_{DB,t}Dummy_{1,t}) \\ &+ \varrho Exp_{j,t} + \varphi Exp_j Dummy_{1,t} \\ &+ \theta Exp_j log(Y_{DB,t}Dummy_{1,t}), \end{aligned} \tag{2}$$

where the dependent variable $Y_{i,j,t}$ is the daily YTM of senior bond i issued by bank j at time t. This variable proxies the bank's solvency risk as perceived by the market. The regressors are (i) the dependent variable lagged by one period to account for serial correlation, $Y_{i,j,t-1}$; (ii) the daily average YTM of the CoCos issued by DB taken in logarithm to account for possible nonlinearities in the estimated relation, $log(Y_{DB,t})$; (iii) the dummy $Dummy_{1,t}$ which identifies the first distress event; (iv) the interaction of $Dummy_{1,t}$ with $log(Y_{DB,t})$; (v) Exp_j to control for banks' exposure to CoCo-specific contagion; (vi) the interaction of the exposure variable Exp_j with $Dummy_{1,t}$; and (vii) the triple interaction between Exp_j, $log(Y_{DB,t})$ and $Dummy_{1,t}$.

Our focus is on the coefficients of $Dummy_{1,t}$ and of the three interactions $log(Y_{DB,t}Dummy_{1,t})$, $Dummy_{1,t}Exp_j$, and $log(Y_{DB,t}Dummy_{1,t}Exp_j)$: (i) a positive coefficient of $Dummy_{1,t}$ would indicate a negative effect of the stress of DB on other banks' solvency; (ii) a positive coefficient of $log(Y_{DB,t}Dummy_{1,t})$

would be evidence of a CoCo-specific contagion on the banks' solvency; (iii) positive coefficients of $Dummy_{1,t}Exp_j$ and $log(Y_{DB,t}Dummy_{1,t}Exp_j)$ would instead indicate that the negative repercussions would have been more severe for the banks more exposed to DB through the CoCo-specific contagion channel.

The results from the estimation of equation (2) are shown in table 11, with the GLS estimation reported in column 1, the FGLS in column 2, and the system GMM in column 3.

The findings across the three different estimations show that around the first distress event there are negative implications from the stress of DB on the financial condition of other banks, which go beyond the documented contagion that occurred within the CoCos market. This is indicated by the positive coefficient of $Dummy_{1,t}$. At the same time, these negative implications have been augmented by the deterioration in the conditions of DB CoCos, thereby indicating that the CoCo-specific contagion has affected not only other banks' CoCos but also banks' senior bonds, as suggested by the positive coefficient of $log(Y_{DB,t}Dummy_{1,t})$. There has not been however a differential impact of the distress of DB and of the CoCo-specific contagion for those banks with a higher exposure to DB, as shown by the nonsignificant coefficients of $Dummy_{1,t}Exp_j$ and $log(Y_{DB,t}Dummy_{1,t}Exp_j)$.

8. Robustness

In this section we conduct two robustness exercises. First we deepen the analysis about the possible role of nonlinearities, and second we check the validity of our results against the existence of omitted variables.

8.1 Nonlinearities in CoCo Prices during Stress Events

As an additional robustness exercise, we test the validity of our results against nonlinearities between bank risk and the price of CoCos during the two DB distress periods, in addition to what is already captured by the log-linear specification used. We do so by estimating equation (1) augmented by the interactions $X_{k,j,t}Dummy_{s,t}$, that is, between the two dummy events and, respectively, the senior bond YTM, the stock return, and the EDF

Table 11. Implications of CoCo-Specific Contagion for Bank Enterprise Value

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Lag Senior Bond YTM	0.970***	1.000***	1.006***
	(0.007)	(0.000)	(0.007)
Log CoCo DB YTM	−8.547***	0.555***	3.364*
	(2.332)	(0.086)	(2.016)
Event1	89.229***	40.824***	23.242
	(28.663)	(5.163)	(31.935)
Event1 × Log(CoCo DB)	33.971***	15.888***	9.167
	(11.272)	(2.022)	(12.484)
Exp. to DB	−0.770	0.064	−0.674
	(0.469)	(0.053)	(0.484)
Exp. to DB × Event1	−9.839	−19.193	54.288
	(68.222)	(12.314)	(58.917)
Exp. to DB × Event1 × Log(CoCo DB)	−2.977	−7.318	21.410
	(26.887)	(4.825)	(23.202)
Constant	−19.399***	1.318***	8.449*
	(5.364)	(0.235)	(4.675)
Observations	124,776	124,776	124,776
Number of ISIN	346	346	346
Hansen Test			266.5
p-value			0
Arellano-Bond Test for AR(1)			−3.393
p-value			0.000690
Arellano-Bond Test for AR(2)			−0.762
p-value			0.446
Number of Instruments			50

Notes: Estimates of the yield to maturity of the CoCos issued by European banks, excluding DB. Random-effect GLS estimation with clustered standard errors by "bank*time" and intragroup correlation is shown in column 1; feasible GLS estimation with heteroskedastic error structure across panels and panel-specific AR(1) process within panels is reported in column 2; Blundell-Bond system GMM estimate is shown in column 3. For the GMM, Hansen, AR(1), and AR(2) tests are provided. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

change of bank j. The results in table 12 confirm the validity of our findings, as the significance of the coefficients of $Y_{DB,t}Dummy_{s,t}$ is unchanged relative to table 6. They also indicate that indeed there are nonlinearities in the relationship between the bank risk and the price of its CoCos during the first DB distress period, as confirmed

Table 12. Fundamental and CoCo-Specific Dependence of CoCos YTM during DB Events: Nonlinearities in the Price of Bank Risk

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Lag CoCo YTM	0.994***	0.997***	0.993***
	(0.002)	(0.000)	(0.001)
Senior Bond YTM	0.158**	0.040***	−0.027
	(0.073)	(0.012)	(0.201)
Stock Return	−0.079***	−0.086***	−0.228***
	(0.011)	(0.002)	(0.051)
EDF Change	0.800**	0.159	−2.366
	(0.315)	(0.121)	(3.906)
Distance to Trigger	−0.023**	−0.010***	−0.008
	(0.012)	(0.003)	(0.012)
CoCo DB	0.035	0.054***	−0.367
	(0.029)	(0.008)	(0.280)
CoCo DB × Event1	1.165***	1.399***	3.455***
	(0.208)	(0.048)	(0.825)
CoCo DB × Event2	−0.057	−0.078	0.010
	(0.123)	(0.082)	(0.255)
Event1	−0.092***	−0.108***	−0.266***
	(0.014)	(0.004)	(0.063)
Event2	0.004	0.005	−0.002
	(0.009)	(0.006)	(0.018)
Senior Bond YTM × Event1	0.160	0.054	−0.001
	(0.194)	(0.039)	(0.132)
Senior Bond YTM × Event2	0.170	0.137	0.242
	(0.128)	(0.103)	(0.301)
Stock Return × Event1	−0.076	−0.014*	0.358***
	(0.072)	(0.008)	(0.115)
Stock Return × Event2	−0.046	−0.024	−0.111
	(0.032)	(0.021)	(0.078)
EDF Change × Event1	5.491*	5.173***	16.751**
	(2.924)	(0.549)	(8.035)
EDF Change × Event2	3.378	1.511	10.627
	(2.531)	(1.649)	(7.096)
CoCos Characteristics:			
Subordination Type	−0.004*	−0.001***	−0.005***
	(0.002)	(0.000)	(0.001)
Equity Conversion	−0.000	0.000	−0.001
	(0.001)	(0.000)	(0.001)
Permanent Write-Down	0.002*	0.001**	0.001
	(0.001)	(0.000)	(0.001)

(*continued*)

Table 12. (Continued)

Variables	GLS RE (1)	FGLS (2)	System GMM (3)
Cancellation by Regulators	-0.001^{*}	-0.000^{**}	-0.001^{**}
	(0.001)	(0.000)	(0.001)
Trigger Level	0.001	0.000^{**}	-0.001
	(0.001)	(0.000)	(0.001)
Size	0.000	-0.000	0.000
	(0.001)	(0.000)	(0.001)
Constant	-0.019^{***}	-0.011^{***}	0.009
	(0.005)	(0.001)	(0.021)
Observations	37,987	37,987	37,987
Number of ISIN	82	82	82
Hansen Test			78.90
p-value			0.270
Arellano-Bond Test for AR(1)			-1.837
p-value			0.066
Arellano-Bond Test for AR(2)			0.337
p-value			0.736
Number of Instruments			95

Notes: Estimates of the yield to maturity of the CoCos issued by European banks, excluding DB. Random-effect GLS estimation with clustered standard errors by "bank*time" and intragroup correlation is shown in column 1; feasible GLS estimation with heteroskedastic error structure across panels and panel-specific AR(1) process within panels is reported in column 2; Blundell-Bond system GMM estimate is shown in column 3. For the GMM, Hansen, AR(1), and AR(2) tests are provided. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

by the significant coefficients of both variables $Stock_{j,t}Dummy_{s,t}$ (columns 2 and 3) and $EDF_{j,t}Dummy_{s,t}$ (columns 1, 2, and 3).

8.2 Omitted Variables

In two out of the three estimations presented in table 6, the DB CoCo variable is not a significant driver of other banks' CoCo YTM under normal market conditions, as we expected. However, in the estimation (column 2) it turned out to be significant. Even though it is so only in one out of three estimations, we suspect that its significance might be due to a possible omitted-variable problem, which

is to say that omitted or unobservable variables might drive both DB's and other banks' CoCos YTM, hence the significance of the DB's CoCos in table 6.

In order to test for this possibility, we perform a four-step procedure which aims to separate the effect of any omitted variable from any DB-specific effect. The procedure is the following. First, we run a principal component analysis (PCA) on the data set of CoCos YTM, which includes both the CoCos issued by DB and those issued by the other banks. By means of the PCA, we aim to identify those components that capture the CoCos co-movements. As such, they should also reflect the effect of any unobserved variable that affect all CoCos similarly. Second, we select the components that together explain a significant proportion of the overall variance of our CoCos data set; in particular, we consider the first four components that together account for about 87 percent of the variance. Third, we use them as regressors to explain the DB's CoCos YTM; the resulting residuals are stored, as they constitute the "purely" idiosyncratic part of the DB's CoCos YTM which remains unexplained from the regression. Fourth, with an approach similar in spirit to the forecasting method proposed by Stock and Watson (2002), we study the CoCo YTM running again the FGLS estimate of equation (1), but instead of including the DB CoCo YTM and its interactions with the two stress events, we use as regressors (i) the first four principal components selected in the previous step; and (ii) the stored DB-specific residuals, which we also interact with the two DB stress events. Following this procedure, we should be able to isolate any true DB-specific effect from that of any omitted or unobservable variable.

The results of the FGLS regression, which are comparable to those presented in table 6, column 2, are shown in table 13. They confirm our hypothesis that the significance of the CoCo DB variable evidenced in table 6, column 2 is more likely to be due to an omitted-variable problem rather than to a DB-specific role in explaining other banks' CoCos YTM in normal times. In fact, the results show that the DB-specific variable (DB residuals) is not significant once we control for possible omitted variables with the first four principal components (all significant). These results also strengthen our argument about the existence of a CoCo-specific contagion channel from DB to other banks during the first stress event, as the interaction of

Table 13. Controlling for Omitted Variables

Variables	**FGLS**
Lag CoCo YTM	0.997***
	(0.001)
Senior Bond YTM	0.064***
	(0.023)
Stock Return	−0.096***
	(0.003)
EDF Change	0.813***
	(0.150)
Distance to Trigger	−0.010**
	(0.004)
Component 1	0.000***
	(0.000)
Component 2	0.000***
	(0.000)
Component 3	−0.001***
	(0.000)
Component 4	0.000***
	(0.000)
DB Residuals	−0.071
	(0.098)
Event1	−0.004***
	(0.000)
Event2	0.000
	(0.000)
DB Residuals × Event1	4.767***
	(0.157)
DB Residuals × Event2	−0.302
	(0.197)
CoCos Characteristics:	
Subordination Type	−0.001**
	(0.000)
Equity Conversion	0.000
	(0.000)
Permanent Write-Down	0.001
	(0.000)
Cancellation by Regulators	−0.000*
	(0.000)

(*continued*)

Table 13. (Continued)

Variables	FGLS
Trigger Level	0.000
	(0.000)
Size	−0.000
	(0.000)
Constant	−0.008***
	(0.001)
Observations	17,788
Number of ISIN	80

Notes: Feasible GLS estimation of the yield to maturity of the CoCos issued by European banks, excluding DB. The estimation allows for heteroskedastic error structure across panels and a panel-specific AR(1) process within panels. Robust standard errors are in parentheses. *** $p < 0.01$, ** $p < 0.05$, * $p < 0.1$.

the DB residuals variable with the first stress-event dummy is very significant.

9. Conclusions and Policy Implications

Our empirical analysis provides evidence of, on the one hand, the existence of a CoCo-specific contagion channel in the propagation of the distress of DB to the rest of the European banks in the first stress episode considered and, on the other hand, the fact that such a contagion channel seems to have waned in the second episode.

From a financial stability perspective, the widespread turbulence in the CoCos market during an idiosyncratic shock is not totally explained by fundamentals and supports the idea of CoCo skeptics that these instruments might prove destabilizing. In fact, in the aftermath of the January–February turmoil, some senior bank executives complained that the rules for CoCos are too complicated and could undermine a bank's financial position rather than strengthen it in a crisis (Arnold and Hale 2016). Concerns regarding these instruments were compounded by the uncertainties surrounding the interaction between the results of the EU-wide stress test and the pillar 2 requirement, together with the direct consequences of this interaction on the level of the MDA trigger point, i.e., the capital level

below which a bank faces restrictions on the amount of distributable profits, including in the form of coupon payments on CoCos.

The ECB (2016b) clarified its position on July 29, 2016, in apparent coordination with the simultaneous publication of the stress-test results by the EBA (2016a).[23] In particular, the ECB not only reduced uncertainty on the supervisory framework in a dimension that is crucial for investors to understand the repayment behavior of CoCos, but also de facto increased the protection of CoCos' investors by making it less likely that banks would face restrictions on the payment of coupons on AT1 instruments.[24]

Both the reduction in uncertainty and the enhancement of protection are likely to have played a role in the eliminating the CoCo-specific contagion in the second DB stress period identified in our empirical analysis. In addition to that, the absence of Coco-specific contagion in the second DB stress period could also result from a "learning process" from investors that made agents more aware of the riskiness of these instruments at the time of the second stress event, thereby rendering CoCos' prices less volatile during idiosyncratic shocks. More market stability in stressed periods is certainly reassuring regarding the convenience of CoCos as a substitute of costlier common equity for the purpose of increasing banks' loss-absorption capacity.

We conclude the paper with a final comment. One of the most appraised features of CoCos is that of enabling the recapitalization of an institution as a going concern, that is, ahead of resolution. This property distinguishes CoCos from other AT2 instruments such as subordinated debt or bail-in eligible liabilities (Minimum Requirement for own funds and Eligible Liabilities, MREL) that allow for the bank recapitalization as a gone concern, that is, in a resolution

[23]It announced that one element of the quantitative results of the stress test (that is the potential capital shortfall under adverse scenario) would be an input of the nonbinding pillar 2 guidance, a newly introduced element of the Supervisory Review and Evaluation Process (SREP) that, as opposed to pillar 2 requirements, does not contribute to the determination of the MDA trigger.

[24]In the European Central Bank's (2016a) own words, with "the introduction of the component of Pillar 2 guidance, the capital requirements of a bank in terms of Pillar 1 plus Pillar 2 requirements will be reduced – all things being equal. As a result, the trigger for the maximum distributable amount (MDA) will go down — also all things being equal."

process. Yet, the supervisory actions described above went in the direction of relaxing the conditions that may lead to automatic restrictions on the payment of CoCos' coupons, thereby reducing the ability of these liabilities to provide for additional capital as a going concern, and rendering them de facto closer to AT2 instruments. While the ECB actions might have been justified (and the need for a reduction on uncertainty on supervisory treatment was uncontroversial), neither supervisory authorities nor investors should interpret them as an implicit recognition that "whatever it takes" will be done to stabilize CoCos markets. These instruments will only be useful if they achieve in practice what they are supposed to do by their design, namely the recapitalization of an institution as a going concern in a smooth manner that disrupts neither the institution itself nor the rest of the system.

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Risk Shocks and Monetary Policy in the New Normal*

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Risk shocks give rise to a tradeoff for monetary policy between inflation and output stabilization in the canonical New Keynesian model if they are large relative to the distance between the nominal interest rate and its lower bound. The tradeoff-inducing effects operate through expectational responses to the interaction between the perceived volatility of conventional level shocks and the available monetary policy space. At the same time, a given monetary policy stance becomes less effective. Optimal time-consistent monetary policy therefore calls for potentially sharp cuts in interest rates when risk is perceived to be elevated, even if this risk does not materialize in any actual disturbances to the economy. The new normal for monetary policy may be one in which policymakers should both constantly lean against a tendency for inflation expectations to anchor below target—operating the economy above potential in the absence of disturbances—whilst accepting that inflation will settle potentially materially below target, and respond nimbly to changes in public perceptions of economic risk.

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1. Introduction

During the Great Moderation, there was a general consensus that spells at the zero lower bound (ZLB) would be rare and short (see, e.g., Coenen, Orphanides, and Wieland 2004, Reifschneider and Williams 2000, and Schmitt-Grohé and Uribe 2010). Over the past decade, this pre-crisis consensus has been revised in light of the incoming data (e.g., Blanchard 2014, Chung et al. 2012, Kiley and Roberts 2017, and Williams 2014). Equilibrium real interest rates are now widely expected to recover only to levels that fall short of historical averages—reducing the scope for future cuts in policy interest rates—and disturbances are expected to be larger—increasing the occasional need for such cuts. In the future, policy rates are deemed likely to hit their lower bounds more frequently than previously assumed. Nevertheless, an optimistic view holds that unconventional monetary policies such as quantitative easing and Odyssean forward guidance can be relied upon as substitutes for conventional reductions in policy rates (e.g., Bernanke 2017, Harrison 2017, Kiley 2018, and Reifschneider 2016). Whilst the ZLB may bind from time to time, monetary policy's extended toolkit will rarely be constrained according to this view.[1] But suppose the public are not fully convinced by such assurances. Suppose the economy recovers to a "new normal" (in the terminology of El-Erian 2010) in which people occasionally find reason to worry that policymakers may not be able to respond to future adverse disturbances with sufficient monetary stimulus. How should monetary policy be conducted in such an environment? Are the prescriptions for good monetary policy in "normal times" developed during the Great Moderation sufficient guideposts for determining the appropriate stance of policy?

This paper points to two key differences that set monetary policy in the "new normal" apart from that of the greatly moderated pre-crisis economy. First, when risk is high relative to the available

[1]On a similar note, the view that the Great Recession marked the end of the Great Moderation is disputed; see, for example, Gadea, Gómez-Loscos, and Pérez-Quirós (2018). But even in their analysis, a continuation of the Great Moderation mainly manifests itself in a slow recovery from the Great Recession.

monetary policy space, policymakers should operate the economy above its efficient potential in normal times. This stimulatory bias leans against a tendency for inflation expectations to re-anchor too far below target, but it does allow inflation to settle potentially materially below target in the absence of disturbances. Welfare may be improved by appointing an independent central banker with a slightly higher inflation target then the social optimum. Second, because of constraints on monetary policy alone, changes in the public's perception of risk affect the appropriate stance of monetary policy through time-variation in the appropriate tradeoff between inflation and real stability. A spike in uncertainty, for example, has a negative cost-push effect because of the ZLB alone and makes a given stance of monetary policy less effective. Potentially sizable changes in interest rates may be warranted even if risk does not materialize in any actual disturbance to the economy. This is in sharp contrast to conventional guidelines derived in the optimal monetary literature under conventional perfect foresight assumptions and without considering the ZLB, in which (in the absence of precautionary behavior by households and firms) the appropriate stance of monetary policy is affected only by shocks that have actually occurred or are fully anticipated; see, e.g., Clarida, Galí, and Gertler (1999) and Svensson and Tetlow (2005).

I derive these results in a simple version of the canonical New Keynesian model. The monetary policy design problem for this model served as the "science of monetary policy" for the Great Moderation (Clarida, Galí, and Gertler 1999), and it remains the theoretical foundation for the kind of flexible inflation targeting effectively practiced by major central banks today (Svensson 2010). Throughout, I focus on responses under optimal time-consistent monetary policy. The period-by-period nature of decisionmaking under discretion makes it a realistic description of the actual conduct of monetary policy in a flexible inflation-targeting regime (see, e.g., Bean 2013). In particular, policymakers set interest rates policy meeting by policy meeting to achieve good outcomes given their operational targets. Neither do they follow an instrument rule mechanically, nor do they commit both themselves and future incumbents to a policy plan that will later turn out to be undesirable. In line with conventional wisdom, policymakers cannot bootstrap the economy out of a ZLB episode by promising a future economic boom as advocated by

Eggertsson and Woodford (2003).[2] As noted by Kiley (2018), it is doubtful that any central bank has attempted such purely Odyssean forward guidance in response to binding lower bounds on short-term policy interest rates.

Specifically, I solve for the risky steady state and I study optimal responses to risk shocks around that steady state in a quasi-linear version of the New Keynesian model augmented with a ZLB. In line with the definition in Coeurdacier, Rey, and Winant (2011), the risky steady state is the point at which the economy settles when previous shocks have abated but agents are aware that further shocks may hit in the future. The *risky* steady state differs from the perfect-foresight or *deterministic* steady state, in which agents do not consider the possibility of future shocks. By risk shocks I mean changes in the standard deviations of conventional level shocks in the model. I trace out responses to such changes along the zero-shock path, i.e., the trajectory of the economy over time in which innovations to level shocks do not actually occur. As it were, nothing actually happens in this paper. But crucially, agents remain aware that level shocks could hit the economy at any time in the future and take the economy off the zero-shock path. For analytical tractability, I maintain the assumption in my baseline analysis that agents form expectations at any given point in time in the belief that current risk levels will persist. With this assumption, the effects of risk shocks can be thought of as the economy's responses to changes in a broad notion of the public's perception of risk. A robustness exercise shows how results generalize to the case where agents form fully rational expectations about a stochastic risk shock process as well as about the level shock processes.

I refer to these changes in second moments as risk shocks following the traditional Knightian distinction between risk and uncertainty.[3] Agents in the model economy have well-defined probability

[2]Adam and Billi (2006) show that risk increases welfare gains from time-inconsistent policy plans if policymakers could find ways to credibly commit to them.

[3]While the macroeconomic literature often follow Bloom (2009) in referring to second-moment shocks as uncertainty shocks, this term is also used to describe a range of related phenomena; see, e.g., Kozeniauskas, Orlik, and Veldkamp (2018) for a recent discussion. Fernández-Villaverde et al. (2011) also refer to second-moment shocks as changes in risk, and LeRoy and Singell (1987) discuss the Knightian terminology.

distributions of economic shocks in mind when making decisions, and at any given point in time these distributions coincide with the actual ones. Probabilities are only subjective in the limited sense that agents do not foresee future changes in the levels of standard deviations. We can think of changes in these probabilities as changes in perceived risk only because there is an observational equivalence between the effects of changes in actual and subjective probabilities along the zero-shock path. Elevated risk in the model may well stand in for Knightian uncertainty in reality, but I make no attempt to model such uncertainty about probability distributions directly as, for example, in the work by Ilut and Schneider (2014) and Masolo and Monti (2017).

It is important to note that the macroeconomic risk shocks considered here are very different from the cross-sectional risk shocks analyzed by Christiano, Motto, and Rostagno (2014).[4] In their paper, a "risk shock" refers to a disturbance to the ex post realization of the dispersion of the quality of capital acquired by entrepreneurs. When this dispersion widens, the agency problem associated with financial intermediation becomes more severe. As credit spreads increase, entrepreneurs demand less capital and aggregate demand contracts for a given stance of policy. In the simple New Keynesian model I consider, such a scenario would correspond to a negative *level* shock to the efficient-equilibrium real rate of interest. To keep aggregate demand in line with the economy's supply potential, monetary policy would have to counter cross-sectional risk shocks with a looser policy stance.

The simple quasi-linear structure I consider comes with the benefit that risk affects the economy exclusively through its interaction with the ZLB. Schmitt-Grohé and Uribe (2004) show that risk does not affect decision rules to a first-order approximation, and only the constant term up to a second order, for a general class of models *without* inequality constraints. This class of models includes the textbook New Keynesian model without a ZLB. In my analysis of the New Keynesian model, risk affects decision rules to a first order because of the additional inequality constraint on policy rates—and

[4]Similarly, they are different from the volatility fluctuations considered by Arellano, Bai, and Kehoe (2019).

for that reason only. Second-moment shocks have real effects in the model through expectations if and only if the ZLB is binding in some conceivable states of the world. This allows me to focus on the defining feature of the new environment—an expected regular recurrence of a binding ZLB—and its implications for monetary policy in normal times, without conflating them with effects of risk not stemming directly from a potentially binding ZLB.

I solve the model following the stochastic extended path approach in Evans et al. (2015). On the one hand, the approach follows Adjemian and Juillard (2013) in relaxing the assumption of certainty equivalence maintained in the extended path procedure first proposed by Fair and Taylor (1983). On the other hand, a number of assumptions are imposed to simplify the analysis and focus on the expectational effects of risk directly caused by the ZLB.[5] Compared with a full global solution procedure, the approach relies on (i) a first-order approximation of behavioral relations with the ZLB imposed as the only nonlinearity (through which risks have first-order effects); (ii) the absence of intrinsic persistence (which could occur, for example, because of habits in consumption, indexation in pricing, or capital accumulation); (iii) optimal monetary policy under discretion (so that policymakers do not seek to affect expectations through commitments); and (iv) the approximation of continuous autoregressive shock processes as discrete-space Markov chains. As these assumptions combine to eliminate state variables, the model can be solved in a simple recursive procedure, providing a mapping between risk and macroeconomic outcomes.

The paper is organized as follows. First, section 2 discusses the relation to the literature. Section 3 then describes the model, section 4 the solution method, and section 5 the calibration. Section 6 presents the risky steady state, and section 7 shows responses to risk shocks. Section 8 presents normalization scenarios after a ZLB episode, and section 9 presents an application with stochastic volatility. Finally, section 10 concludes.

[5]These simplifications avoid the curse of dimensionality that forces Adjemian and Juillard (2013) to prune the tree of forward histories when calculating expectations. See Maliar et al. (2015) for a discussion as well as a generalized procedure.

2. Relation to the Literature

The analysis follows previous studies of the implications for discretionary monetary policy of deviations from certainty equivalence in New Keynesian models with a ZLB. Adam and Billi (2007) and Nakov (2008) first showed how the interaction of risk and the ZLB may give rise to a negative bias in expectations in the New Keynesian model's stochastic equilibrium. Both papers illustrate how this bias amplifies the economy's responses to negative shocks to the level of the equilibrium real rate of interest and leads to tradeoffs when monetary policy is driven close to the ZLB. They show how gains from commitment are significantly larger as a result when risk is taken into account. In a recent application, Nakata and Schmidt (2014) further suggest that the skew in expectations provides a justification for a weight-conservative central banker. Similarly, Evans et al. (2015) emphasize that higher risk generally calls for looser monetary policy when the ZLB may bind. They find that liftoff from a ZLB episode should be delayed when agents are concerned about the risk of future episodes.

Compared with these seminal papers, my contribution is, first, to characterize the economy's risky steady state when monetary policy's room for maneuver may be deemed inadequate in normal times. Writing during the "old normal" when ZLB episodes were expected to be rare, Adam and Billi (2007) and Nakov (2008) did not pay close attention to steady-state outcomes or distinguish clearly between different notions of the steady-state and average outcomes. Instead, they emphasized the amplification of effects following a large persistent fall in the equilibrium real rate of interest from its normal level.[6] Second, I trace out dynamics around the risky steady state following changes in risk, illustrating how a risk shock propagates through dynamics in expectations in a way that induces time-varying tradeoffs for monetary policy. I show that this time-variation calls for monetary policy responses even in the absence of actual disturbances (such as level shocks to the equilibrium real rate), and I show that these responses do not depend on the source of risk (i.e., the

[6] Also, Evans et al. (2015) operate in a part of the parameter space for which the solution to the New Keynesian model with a ZLB is explosive, pinned down only by the assumed expectational horizon.

particular shock process for which risk has changed). Moreover, I show that impulse responses to positive and negative risk shocks are asymmetric around the risky steady state because of a nonlinearity in the mapping between risk and economic outcomes.

The simple quasi-linear framework sets my analysis apart from the recent related work by Basu and Bundick (2015, 2017). In these papers, second-moment shocks have real effects through precautionary savings behavior by households. The authors show that these higher-order effects are greatly amplified at the ZLB as monetary policy fails to respond to them.[7] Moreover, a feedback mechanism sets in: monetary policy's inability to respond to what amounts to a further fall in the equilibrium real rate reduces the expected mean of outcomes, in turn inducing further precautionary saving.[8] In my analysis, by contrast, risk shocks affect the economy solely because of changes in expected mean paths for macroeconomic variables for a given equilibrium real rate of interest. The quasi-linearity thus serves both to illustrate that risk shocks may have significant real effects through the ZLB without precautionary behavior and to isolate such direct effects from those effectively operating through shifts in the equilibrium real rate.[9] There are two other significant differences. First, Basu and Bundick (2015, 2017) either impose a simple instrument rule for monetary policy or allow for commitment policies that essentially assume away the effect of the ZLB on mean expectations. Instead, I consider policy prescriptions under the more realistic assumption that optimizing policymakers do not have access to such commitment devices. Second, Basu and Bundick (2015, 2017) are concerned with the contributions of second-moment shocks to macroeconomic volatility over the past. Specifically, they

[7]Basu and Bundick (2015, 2017) compute third-order approximations to behavioral relations to capture effects operating through precautionary behavior when monetary policy is unconstrained, while they turn to global solution methods when imposing the ZLB.

[8]See also Bloom (2009) and Nakata (2017) for discussions of how other nonlinearities may give rise to real effects from movements in risk that may be amplified by the ZLB. Fernández-Villaverde et al. (2015) and Johannsen (2014) suggest that uncertainty about fiscal policy in particular has larger implications when monetary policy is constrained.

[9]See Paoli and Zabczyk (2013) for an analysis of the effect of precautionary saving on the level of the effective equilibrium real rate of interest.

explain how elevated macroeconomic risk could have contributed significantly to the Great Recession, while changes in second moments amounted to inconsequential background noise before the financial crisis. Hence, they distinguish sharply between periods when the ZLB is both nonbinding and a negligible risk, and periods when the ZLB is in fact binding. By contrast, I derive practical normative prescriptions for monetary policy in an environment in which the ZLB is not necessarily binding but nevertheless constantly looming as a cloud on the horizon.

My paper is also closely related to the contemporaneous and independent work by Hills, Nakata, and Schmidt (2016). In their paper, the authors also compare the risky and deterministic steady states in a stylized New Keynesian model, finding results fully consistent with mine. But they then go on to quantify the difference between these steady states in a richer model calibrated to match key features of the U.S. data over the past decades. I instead focus on the risky steady state in a hypothetical new normal, and I proceed to consider the implications of potential time-variation in the underlying level of macroeconomic risk. Moreover, while Hills, Nakata, and Schmidt (2016) consider outcomes under a simple monetary policy rule taken to be representative of monetary policy's reaction pattern in the past, I derive normative prescriptions by solving for optimal monetary policy responses for the future. Finally, while they solve their model in nonlinear form using global methods, I keep my analysis within a quasi-linear framework to facilitate a direct comparison of these prescriptions with those of the influential "science of monetary policy" (Clarida, Galí, and Gertler 1999).[10] I thus take the two papers to be complementary, each providing a different perspective on the interactions between risk and the ZLB.

More broadly, my paper relates to a growing literature on the effects of risk and uncertainty. Following the work of Bloom (2009), there has been a surge of interest in this issue. Whilst the empirical literature has struggled to identify structural risk shocks from the volatility and forecast disagreement measures that are usually taken to be proxies for risk and uncertainty, theoretical work has provided clear channels through which risk shocks may affect the

[10]See, e.g., Carney (2017) for an application to monetary policy in practice.

economy; see, e.g., the survey by Bloom (2014). The expectational mechanism at the core of my analysis is a "bad news channel" in the terminology of Bernanke (1983). It arises because monetary policy is sometimes unable to provide sufficient stimulus in response to large adverse shocks, while it can always act to contract the economy when needed. The model's prediction that the effect of a given risk shock is larger the closer the economy is to the ZLB is in line with the empirical evidence provided by Caggiano, Castelnuevo, and Pellegrino (2017) and Plante, Richter, and Throckmorton (2018). Similarly, the finding that monetary policy is less effective when risk is high is consistent with the evidence in Aastveit, Natvik, and Sola (2017). Finally, Caggiano, Castelnuevo, and Nodari (2018) provide recent evidence that monetary in the United States has indeed responded to changes in aggregate risk.

3. The Model

The model is the canonical New Keynesian model, expressed in log-deviations from its deterministic steady state, extended with a ZLB on interest rates. In addition to a specification of monetary policy, it consists of the following equations:

$$\pi_t = \beta E_t \pi_{t+1} + \kappa x_t + u_t \tag{1}$$

$$x_t = E_t x_{t+1} - \frac{1}{\varsigma}\left(i_t - E_t \pi_{t+1} - r_t^*\right) \tag{2}$$

$$i_t + i^* \geq 0, \tag{3}$$

where E_t is the expectations operator, π_t is inflation at time t in deviation from its target π^*, x_t is the output gap defined as output in deviation from its efficient level, and i_t is the nominal interest rate in deviation from its normal deterministic steady-state value i^*. The first equation is the New Keynesian Phillips curve, the second is the forward-looking IS curve, and the third imposes the ZLB. The model is derived from its microfoundations by Galí (2008) and Woodford (2003) among others.

There are two shock processes in the model. The term u_t is a cost-push process, and r_t^* is the efficient-equilibrium real interest rate in deviation from its steady-state level $r^* \approx i^* - \pi^*$. I assume that

the latter is the sum of a deterministic but potentially time-varying component ρ_t and a stochastic process ϵ_t so that $r_t^* = \rho_t + \epsilon_t$. Both the stochastic component of the equilibrium real interest rate and the cost-push shock are given as first-order autoregressive processes with zero-mean Gaussian innovations:

$$\epsilon_t = \mu_\epsilon \epsilon_{t-1} + \nu_{\epsilon,t} \tag{4}$$

$$u_t = \mu_u u_{t-1} + \nu_{u,t}, \tag{5}$$

where $\nu_{\epsilon,t} \sim N(0, \sigma_{\epsilon,t}^2)$ and $\nu_{u,t} \sim N(0, \sigma_{u,t}^2)$. I allow the standard deviations of the innovations to vary over time as indicated by the time subscripts in $\sigma_{\epsilon,t}$ and $\sigma_{u,t}$.

I define a risk shock as a change in one or both of these standard deviations. The standard deviations are independent of each other and can also change independently. As a baseline, however, I consider the special case in which $\varsigma^{-1}\sigma_{\epsilon,t} = \sigma_{u,t} = \sigma_t$ with

$$\sigma_t = \bar{\sigma} + \mu_\sigma \left(\sigma_{t-1} - \bar{\sigma}\right) + \nu_{\sigma,t}, \tag{6}$$

where $\nu_{\sigma,t}$ is an innovation to risk, and $\bar{\sigma}$ is an underlying level of risk in the absence of risk shocks.[11]

Under optimal policy under discretion, a policymaker, hypothetically unconstrained by the ZLB in (3), minimizes the period loss function

$$L \propto \pi_t^2 + \lambda x_t^2 \tag{7}$$

each period subject to the Phillips curve in (1) while taking expectations as given. The loss function can be derived as a quadratic approximation of the utility of the representative household in the full New Keynesian model (again, see, e.g., Galí 2008 and Woodford 2003). The optimality condition takes the form of a conventional targeting rule,

$$\pi_t = -\frac{\lambda}{\kappa} x_t, \tag{8}$$

[11] I only consider realizations of risk that are strictly larger than zero. More broadly, the risk shock process may be specified in logs to rule out nonzero realizations.

stating the optimal policy tradeoff between inflation and the output gap. Following tradeoff-inducing shocks, monetary policy seeks to keep deviations of inflation from target and of output from potential of opposite signs, letting inflation absorb more—and output less—of the adjustment the higher the weight on the output gap in the loss function and the flatter the Phillips curve (i.e., the higher the sacrifice ratio). The interest rate consistent with this optimal allocation can now be found from the IS curve in (2). Since the policymaker is, in fact, constrained by (3), the targeting rule in (8) is replaced by the Kuhn-Tucker conditions

$$(i_t + i^*)(\lambda x_t + \kappa \pi_t) = 0 \tag{9}$$

$$i_t + i^* \geq 0 \tag{10}$$

$$\lambda x_t + \kappa \pi_t \leq 0. \tag{11}$$

As shown formally in appendix A, these conditions imply that the interest rate will be set to the maximum of the level consistent with satisfying the targeting rule in (8) and zero.

4. Solution Method

In the main analysis, I assume that agents observe the current level of risk without expecting further changes in this level to occur—even if risk may actually change over time. This assumption simplifies the calculation of the model solution considerably without affecting results qualitatively. In this section, I outline the solution method as well as the definitions of the stochastic steady state and of a simple impulse response function under this simplifying assumption. In section 9, I discuss and present a generalization to the case with stochastic volatility where agents understand the stochastic nature of the risk shock process as well as the processes for the level shocks.

4.1 Solution

I solve the quasi-linear version of the canonical model following the approach in Evans et al. (2015). I approximate the shock processes by independent Markov processes using the Rouwenhorst (1995) algorithm provided by Galindev and Lkhagvasuren (2010). I then solve the model backwards from a distant future period T, beyond

which there is no risk and all shocks are zero so that $E_t\pi_{t+1} = E_t x_{t+1} = 0$ for all $t > T$. In each step, I take expectations as given and calculate the unconstrained outcome under each policy regime for a state grid of values for the shock processes, (ϵ, u). I then check if this outcome is consistent with the ZLB in (3) for each node in the grid. If so, I take the unconstrained outcome as the solution for this particular node. If not, I calculate the outcome from the model equations with $i_t = -i^*$ imposed. I then update the ex ante expectations of inflation and the output gap using the Markov transition matrices before progressing to the previous period. The solution consists of the $n_\epsilon \times n_\epsilon$ matrices for inflation, the output gap, and the interest rate, to which this algorithm converges in the initial period $t = 0$. See appendix B for further details.

4.2 Stochastic Steady State

The values for the nodes $(\epsilon = 0, u = 0)$ represent outcomes in the event that no nonzero shock has actually materialized. This converged zero-shock solution at $t = 0$ represents the risky steady state of the model as defined by Coeurdacier, Rey, and Winant (2011). In particular, let $y_0^{sol}(\epsilon, u \mid \sigma)$ be the state-contingent solution for variable $y_t \in \{i_t, \pi_t, x_t\}$ for a given level of risk. The risky steady state for this variable is then defined as $y_0^{sol}(\epsilon = 0, u = 0 \mid \sigma)$. This is the resting point to which the economy returns when all shocks have dissipated. Nonzero realizations of shocks will of course continuously drive the economy away from this point, and the risky steady state potentially differs from the deterministic one exactly because it accounts for agents' expectations that such deviations will occur. Unconditional expectations are averages weighted by unconditional probabilities over outcomes across the state space. Given the Markov structure, these probabilities can easily be derived from the eigenvectors associated with the unitary eigenvalues of the transition matrices; see, e.g., Ljungqvist and Sargent (2000). Appendix C provides further details.

4.3 A Simple Impulse Response Function

I find simple impulse responses to a risk shock by running a double loop. The outer loop moves forward from period $t = 0$, while the

inner loop solves the model backwards from period T to the period of the current iteration of the outer loop. For each iteration of the outer loop, I reduce the value of σ_ϵ and/or σ_u from an initial spike according to the assumed process. The economy's responses to the risk shock is the sequence of zero-shock solutions found in the outer loop. In particular, the simple impulse response function is defined as

$$\begin{aligned} I^y_{t+n} = y_0^{sol}(\epsilon = 0, u = 0 \mid \sigma_{t+n} = \mu^n_\sigma \sigma') \\ - y_0^{sol}(\epsilon = 0, u = 0 \mid \sigma_{t+n} = \bar{\sigma}) \end{aligned} \tag{12}$$

for $n \in \mathbb{Z}^+$ and where σ' denotes the value of risk on impact of the risk shock in period t.

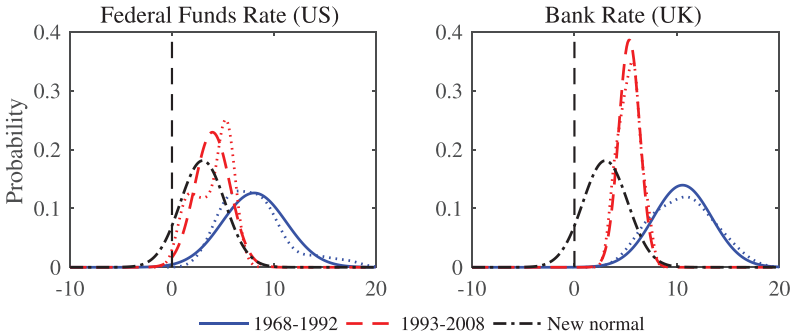
In line with the definition of a traditional impulse response function in Koop, Pesaran, and Potter (1996), this simple function measures the effects of a risk shock hitting the economy at time t on the state of the economy at time $t+n$ given that no other shocks occur. The conceptual experiment is a comparison of the profile for the economy when a risk shock hits with a profile where risk stays at its baseline level, keeping all other shocks (in this case, the level shocks) dormant. But it is a traditional impulse response function with the caveat that agents have myopic rather than rational expectations about the risk shock process itself (while otherwise continuing to form rational expectations). The advantage of taking this approach is that risk does not become a state variable, simplifying calculations considerably. It is also a natural starting point for considering occasional changes in risk perceptions: each period, economic agents simply assign a number to the level of risk they think is present in the economy. But it is a limitation that agents always expect risk to stay constant at a given point in time. As a robustness check, I therefore also show that a generalized specification where agents are allowed to see a stochastic autoregressive profile for risk generates qualitatively similar results (section 9).

5. Calibration

5.1 The New Normal

I calibrate the model to fit a hypothetical "new normal" distribution for desired policy rates—the short-term interest rates a policymaker

Figure 1. Old and New Normal Densities for Policy Rates



Notes: Normal probability density functions with means and standard deviations set equal to sample means and standard deviations for the federal funds rate (left panel) and Bank Rate (right panel) for the sample periods 1968–92 (solid blue lines) and 1993–2008 (dashed red lines) as well as for a hypothetical new normal (dashed-dotted black lines). Dotted lines are estimated kernal distributions with an optimized bandwidth for the normal kernal function for the 1968–92 (in blue) and 1993–2008 (in red) subsamples.

would choose to implement if the ZLB were not a constraint—shown in dashed-dotted black lines in both panels in figure 1. To motivate the choice of this distribution, the figure compares it with recent pre-crisis historical experience in the United States (left panel) and the United Kingdom (right panel). The normal probability density functions shown in solid blue lines share the means and standard deviations with the observed federal funds rate and Bank Rate, respectively, from 1968 through 1992.[12] Blue dotted lines show corresponding kernel density estimates. An observer looking back at these distributions around the time when 2 percent inflation targets were emerging in the early 1990s (see, e.g., Svensson 2010) would not have found much reason to worry about the ZLB. Policy rates had been very volatile over the past quarter of a century (standard deviations were 3.2 percentage points and 2.9 percentage points, respectively), but they had also been high (with means of 8.1 percent and 10.6 percent). The probability that interest rates

[12]For color versions of the figures, see the online version of the paper at http://www.ijcb.org.

should be negative would have seemed negligible. Over the subsequent 15 years, the distributions of policy rates shifted sharply to the left as inflation targeting became established (formally or informally), and inflation expectations anchored at lower levels; see the dashed red lines for normal approximations and dotted red lines for kernel density estimates. In isolation, the lower means (4.0 percent and 5.4 percent) would have increased probability mass below zero. But as volatility fell substantially at the same time, the risk of a binding ZLB did not appear to have increased significantly over the period.

Any Gaussian model informed by interest rate data over the two periods would have produced a probability of negative interest rates very close to zero, including both the semi-structural and the dynamic stochastic general equilibrium (DSGE) models that were increasingly used for such purposes; see, e.g., Chung et al. (2012), Coenen, Orphanides, and Wieland (2004), and Schmitt-Grohé and Uribe (2010). Moreover, assessments of the fit of models estimated using full-information techniques—for example, by comparing fitted normals with kernel distributions (a tougher test than usually applied)—would have pointed to a satisfactory if not perfect fit.[13] It is therefore not surprising that a consensus emerged in the pre-crisis period that the ZLB should be of no great practical concern. As it turned out, however, the estimated models were too sanguine about prospects of a binding ZLB.

Instead of focusing on past distributions, I therefore consider a hypothetical scenario for the normal for monetary policy to which the economy may recover after the financial crisis. Compared with the historical distributions over the 1993–2008 period, it is defined by a further shift to the left and an increase in the spread for desired policy rates. The shift captures an assumed decline in the trend component of the efficient equilibrium real rate of interest. The higher spread reflects an end to the Great Moderation so that

[13]The sample periods in figure 1 are such that estimated distributions are close to normal without detrending. In particular, Bank Rate is very close to normal in each of the two subsamples. The picture is a bit less clear for the United States, where the data point as much to a gradual decline as to a single structural shift in the level around which the federal funds rate fluctuates. The sharp response to the 2001 recession followed by rapid normalization also results in a near-bimodal shape of the estimated kernel for the federal funds rate over the 1993–2008 period.

larger disturbances call for stronger policy responses then in the pre-crisis inflation-targeting period. I set the average desired policy rate to 3 percent—in line with the forecasts emphasized by Reifschneider (2016) and the recent estimates in Del Negro et al. (2017), Kiley (2015), and Laubach and Williams (2016)—and I fix the standard deviation at 2.2, in between the values for the two pre-crisis subsamples for the United States and the United Kingdom. Hence, macroeconomic risk is higher than during the Great Moderation, but not as high as in the preceding decades. With the lower average level of the policy rate, the moderate increase in the desired spread increases probability mass below zero to about 9 percent. In this "new normal" scenario, therefore, the probability that policymakers would set policy rates to negative values if they could is clearly non-negligible.[14]

5.2 Parameterization

In the baseline calibration, I parameterize the model so that it features a distribution for the unconstrained optimal policy rate corresponding to the new normal distribution in figure 1. The (annualized) inflation target is assumed to be $\pi^* = 2\%$ and the deterministic steady-state level of the real interest rate $r^* = 1\%$. The normal nominal interest rate is then approximately $i^* = 3\%$ with a discount factor $\beta = 0.9975$. The deterministic component of r_t^* is set to $\rho_t = 0$ for all t, and the inverse of the elasticity of intertemporal substitution to $\varsigma = 1$ as is common in the literature for this poorly identified parameter (see, e.g., Galí 2008). The slope of the Phillips curve is assumed to be $\kappa = 0.02$. This value is at the lower end of the 0.02–0.05 range of empirical estimates collected by Woodford (2005), in line with the hypothesis that Phillips curves have (if anything) flattened in recent decades (e.g., Blanchard, Cerutti, and Summers 2015).

The weight on the output gap in the loss function is similarly set to $\lambda = 0.02$. This is larger than the $\lambda = \kappa/\zeta$ imposed when the loss function is derived from household utility in the basic

[14]The assumption that desired policy rates are normally distributed is not restrictive. What matters for the qualitative results in the following is that probability mass below zero is non-negligible more than the specific shape of the distribution.

New Keynesian model, where $\zeta > 1$ is the elasticity of substitution between product varieties under monopolistic competition. But for conventional values of ζ around 6 and empirically plausible values of κ, the weight on output would be much smaller than actual mandates for monetary policy seem to imply. The assumed value of λ corresponds to a weight on output stabilization of about a third in annualized terms, a reasonable interpretation of the degree of flexibility in inflation targeting in practice (see, e.g., Carney 2017, English, López-Salido, and Tetlow 2015, and Svensson 2010). Moreover, the basic New Keynesian model is likely to underestimate the appropriate weight on the output gap; see, e.g., Debortoli et al. (2016) and Walsh (2014). With an implied targeting rule for monetary policy with a slope of -1, policymakers seek to let quarterly inflation and the output gap share the burden of adjustments to tradeoff-inducing disturbances equally.

Level shocks are assumed to be moderately persistent with $\mu_u = 0.25$ and $\mu_\epsilon = 0.75$. With these parameter values, an underlying level of risk given by $\sigma = 0.2725/100$ delivers a standard deviation of the unconstrained nominal policy rate of 2.2 when $\varsigma^{-1}\sigma_{\epsilon,t} = \sigma_{u,t} = \sigma_t$ so that the desired policy rate is negative with probability 9 percent. For comparison, I also consider a low-risk scenario with $\sigma = 0.1234/100$. With this low level of underlying risk, the dispersion of desired interest rates is kept low while the mean shifts down to 3 percent. Specifically, the standard deviation of desired interest rate is 1 percentage point, similar to the level observed for the United Kingdom between 1993 and 2008. The probability that interest rate should be negative remains negligible in this case despite a low level of r^*. In section 7, I show impulse responses to risk shocks in this low-risk scenario as well as in the baseline new normal.

Finally, I solve the model with an expectational horizon of $T = 1{,}000$. The solution algorithm converges with significantly fewer iterations than 1,000. Hence, results are not sensitive to this choice of T.

5.3 Model Fit

Table 1 compares the "new normal" baseline and the low-risk scenarios with alternative calibrations in which σ is set to fit the

Table 1. Alternative Calibrations

	Data				Unconstrained Model					
Episode	$E(i)$	$\sigma(i)$	$E(\pi)$	$\sigma(\pi)$	$E(i)$	$\sigma(i)$	$E(\pi)$	$\sigma(\pi)$	100σ	P_0
New Normal	—	—	—	—	3.02	2.20	2.00	2.48	0.27	0.09
Low Risk	—	—	—	—	3.02	1.00	2.00	2.11	0.12	0.00
US 1968–1992	8.07	3.16	5.96	3.73	8.07	3.16	4.16	4.66	0.39	0.01
US 1993–2008	3.97	1.74	2.55	3.59	3.97	1.74	2.00	2.31	0.22	0.01
UK 1968–1992	10.56	2.86	8.77	6.83	10.56	2.86	6.56	6.83	0.35	0.00
UK 1993–2008	5.36	1.03	1.93	2.09	5.36	1.03	2.00	2.11	0.13	0.00

Notes: $E(.)$ and $\sigma(.)$ denote means and standard deviations, respectively, of the nominal interest rate and inflation. Interest rates are measured by the FFR and BR for the United States and United Kingdom, respectively. Inflation is annualized quarterly CPI inflation (source: Datastream). P_0 denotes the probability of negative interest rates in the unconstrained model as well as in normal distributions with means and standard deviations as in the data.

distribution of the unconstrained policy rate to the historical distributions in figure 1. In the historical episodes, I match the mean by assuming that $r^* = 3.75$ in the 1968–92 period, and that $\pi^* = 2\%$ between 1993 and 2016. In all of the alternative calibrations, the probability of negative interest rates is close to zero, rounding to 1 percent only for the United States. Even if the model is too stylized to capture the covariance structure of the data more broadly, the table shows that it gives a fairly good fit to observed inflation volatility across the historical periods. Hence, it does not seem unreasonable a priori that it can shine some light on inflation outcomes in hypothetical scenarios. Nevertheless, I make no claim to provide empirical accuracy in this simple framework. The numerical results should be taken as qualitative indication of the sign and order of outcomes rather than quantitative estimates.

In what follows, I assume that the lower bound is exactly zero as specified in (3). The unconstrained model is hardly a good guide for monetary policy when the desired level of interest rates is negative with a non-negligible probability as in the hypothetical new normal. Absent substantial reform to the payment system, it is improbable that policymakers can persistently drive interest rates into negative territory (e.g., Rogoff 2015). It is more likely that unconventional policies such as quantitative easing may act as substitutes for negative short-term interest rates (e.g., Haldane et al. 2016). The interest rate in the model may best be thought of as a *shadow* rate implicitly incorporating the effects of unconventional policy tools (Black 1995). The shadow rate may be negative if these tools are operational whenever the ZLB binds. But if either the availability or the effectiveness of such tools is limited, the shadow rate will also be bounded from below at some level less than zero. With the ZLB imposed, I effectively assume that cash has not been phased out to allow for negative interest rates, and that unconventional policies cannot act as perfect substitutes for negative interest rates. The analysis becomes irrelevant if either of these assumptions is fully reversed so that interest rates can go negative without difficulty or side effects, or unconventional policy tools can be relied upon as perfect substitutes for changes in policy rates. But the conclusions hold for any combination of the effective lower bound and the distribution of the desired shadow rate such that monetary policy is constrained with a non-negligible probability.

I take such a constellation to be the defining feature of the new normal.

6. Risky Steady State

6.1 Baseline

The first row in table 2 shows risky steady-state outcomes for the interest rate (i), inflation (π), and the output gap (x) in the full model with the ZLB under the baseline calibration. In the new normal, public perceptions of risk are high enough that, because of the ZLB, policymakers are not expected to be able to respond sufficiently to some of the large negative disturbances that are deemed likely to hit the economy in the foreseeable future. By contrast, the public know that monetary policy can always be tightened appropriately, also in response to large inflationary shocks. This asymmetry introduces a negative skew in expectations as first emphasized by Adam and Billi (2007) and Nakov (2008). Under optimal discretionary policy, policymakers lean against the tendency for inflation expectations to anchor below target by operating the economy above potential in normal times through a stimulatory bias in policy rates. Such a long-run effect on real output is feasible because the monetary policy stance interacts with inflation expectations to determine real interest rates in the risky steady state. In this sense, monetary policy is no longer neutral in the new normal's long run. But there are limits to policymakers' willingness to overheat the real economy, and subdued inflation expectations are allowed to weigh on the price setting of firms. Consequently, inflation settles about 20 basis points below target.

Hence, the point at which the economy comes to rest when shocks have faded away does not coincide with the deterministic steady state, in which inflation is on target, the output gap is closed, and the interest rate is at its normal level. Moreover, as table 2 also shows, unconditional expectations deviate from steady-state values. In expectation, output falls short of potential because of spells at the ZLB. These episodes drive the average interest rate above its risky steady-state level, and expected inflation falls somewhat further below target. Notice also that, because of monetary policy's inability to deliver the desired stimulus, the frequency of ZLB

Table 2. New Keynesian Model with ZLB under Optimal Discretion

	Interest Rate			Inflation			Output Gap			
Scenario	i^*	i	$E(i)$	π^*	π	$E(\pi)$	x^*	x	$E(x)$	P_{ZLB}
New Normal	3.02	2.73	2.81	2.00	1.80	1.79	0.00	0.05	−0.01	0.14
r^* Risk Only	3.02	2.94	2.94	2.00	1.93	1.92	0.00	0.02	−0.00	0.07
u Risk Only	3.02	2.98	2.99	2.00	1.98	1.97	0.00	0.01	−0.00	0.04
Lower r^*	2.76	2.25	2.40	2.00	1.66	1.64	0.00	0.09	−0.01	0.20
Lower π^*	2.77	2.27	2.43	1.75	1.43	1.41	0.00	0.08	−0.01	0.20
Higher r^*	3.27	3.10	3.14	2.00	1.88	1.87	0.00	0.03	−0.01	0.10
Higher π^*	3.27	3.09	3.14	2.25	2.12	2.11	0.00	0.03	−0.00	0.10
Very High π^*	5.04	5.03	5.03	4.00	3.99	3.99	0.00	0.00	−0.00	0.01

Notes: i, π, and x are risky steady-state values, stars denote deterministic steady-state values, $E(.)$ and $\sigma(.)$ denote means and standard deviations, respectively, and P_{ZLB} denotes the frequency of a binding ZLB. Deterministic steady-state values for interest rates and inflation satisfy $1 + i^*\% = (1 + r^*\%)(1 + \pi^*\%)$.

episodes is higher than the probability that interest rates should be negative in the unconstrained model. In the new normal, a 9 percent probability that desired policy rates are negative translates into a 14 percent probability that the ZLB is binding.

6.2 Sensitivity to Risk and Policy Space

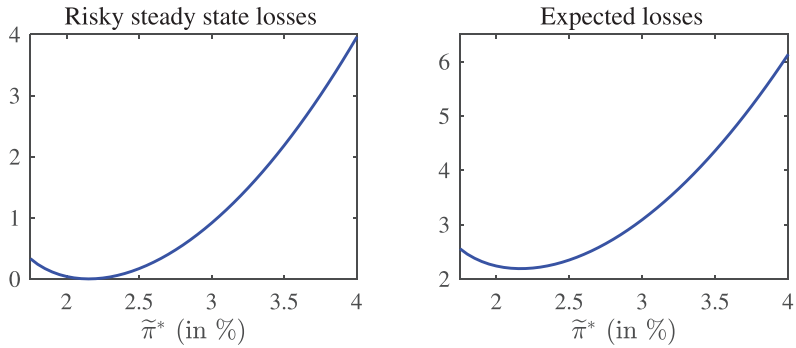
The remaining rows in table 2 illustrate the sensitivity of these statistics to assumptions about risk and the available monetary policy space. In the baseline, I assume that $\varsigma^{-1}\sigma_{\epsilon,t} = \sigma_{u,t} = \sigma_t$ for convenience. But in general, the levels of risk for the two shock processes may not be related by a simple multiple. The second row shows the effect of completely removing the risk of cost-push shocks, while keeping the risk of r^* shocks at the baseline value. The third row shows the opposite case without a perceived risk of r^* shocks. In both cases, the risky steady state deviates from the deterministic one with inflation settling below target. The marginal contributions of the two shocks are similar, but the deviations are much smaller with inflation rates of 1.93 and 1.98, respectively. In the new normal, agents are particularly concerned about the inability of policymakers to respond when large adverse disturbances to the cost-push process and the equilibrium real rate coincide.

Higher risk for individual shocks may, however, result in similar biases as in the benchmark.[15] Inflation may fall short of target in the new normal regardless of the source of risk. Notice also that policymakers have substantial room for maneuver. Inflation falls short of target in normal times only because private agents worry that the ZLB may bind in the future. The re-anchoring of inflation expectations occurs whenever risk is perceived to be high relative to the available monetary policy space.

For a given level of risk, the effect on expectations therefore also depends on the normal distance to the ZLB, as illustrated in the remaining cases in table 2. The closer the economy operates to the ZLB, the larger are the effects of risk on outcomes. If the distance to

[15] Increasing σ_ϵ to about 0.0032 when the risk of cost-push shocks is absent, or σ_u to about 0.0043 when the risk of r^* shocks is negligible, leads to similar biases in inflation and the output gap as under the baseline calibration.

Figure 2. Welfare Losses as a Function of Inflation Target



Notes: Period welfare losses with an optimal annual inflation rate of 2 percent as a function of an operational inflation target denoted by $\tilde{\pi}^*$. The left panel shows losses evaluated in the risky steady state, and the right panel shows unconditionally expected losses.

the ZLB is reduced by about 25 basis points, either because the equilibrium real rate of interest is lower (row 4), or because monetary policy targets a lower inflation rate (row 5), risky steady-state inflation falls below target by a further 15 basis points. When inflation in the deterministic steady state itself is lower, inflation settles almost 40 basis points lower than in the baseline. By contrast, if i^* increases to about 3.25 percent, the negative bias in inflation is reduced by 7–8 basis points. With a higher inflation target, the component of i^* that can be chosen by policymakers, inflation settles around 2.1 percent (row 7). To fully eliminate the negative bias in inflation, however, policymakers will have to target a rate of inflation above 4 percent (row 8).

6.3 Welfare and the Inflation Target

Since inflation is closer to 2 percent and the output gap is closer to zero when monetary policy targets an inflation rate of 2.25 percent (as in row 7 in table 2), welfare losses are unambiguously lower in the risky steady state when evaluated using a loss function (7) that is centered around a 2 percent optimal rate of inflation. In the new normal, welfare may be improved by appointing an independent central banker with a slightly higher operational inflation target than the social optimum. But as shown in figure 2, setting

the operational target too high—for example, at the level which eliminates the negative bias in expectations—comes with substantial welfare costs if the socially optimal rate of inflation is 2 percent. Specifically, only operational targets in the open interval (2.00%, 2.30%) result in smaller welfare losses both in the risky steady state and in expectation, where expected period welfare losses are calculated as the probability-weighted sum of losses across the state space. Provided that monetary policy responds optimally, the simple analysis does not by itself provide a case for increasing the inflation target to, say, 3 percent or 4 percent as suggested as a potential response to low equilibrium interest rates, e.g., by Ball et al. (2016), Blanchard, Dell'arricia, and Mauro (2010), Krugman (2014), and Williams (2009).

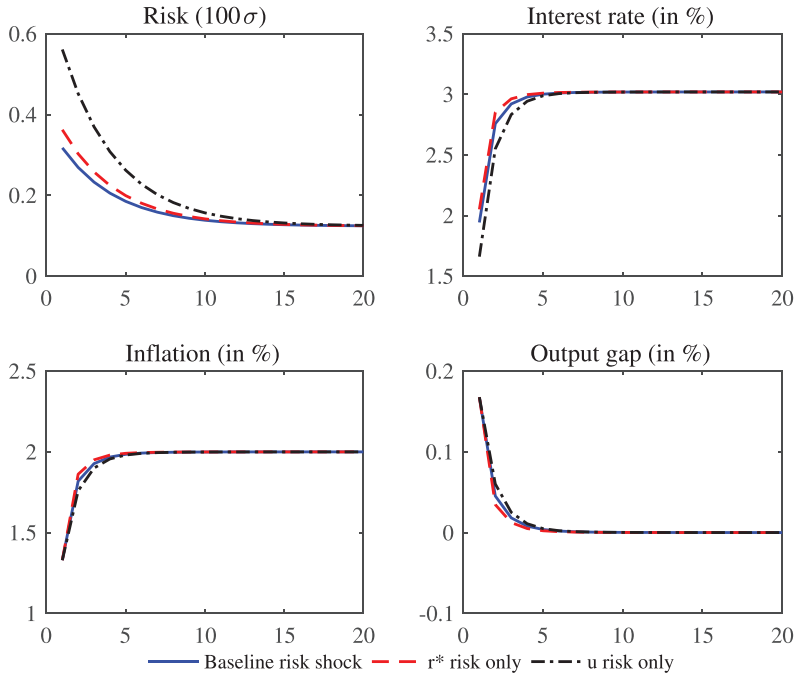
7. Impulse Response to Risk Shocks

Now suppose that risk may vary over time. How does the economy adjust to changes in the perception of risk? To build intuition, I first present simple impulse responses to the baseline risk shock in (6) starting from a low-risk steady state. The low-risk case differs from the new normal only in that $\sigma = 0.0012$. With this low level of underlying risk, the probability that interest rates should be negative remains negligible despite a low level of r^* (see the second row in table 1). The risky steady state therefore practically coincides with the deterministic steady state. I then consider a risk shock for each of the two level shocks in turn, before turning to a baseline risk shock in the risky steady state implied by the new normal scenario.

7.1 The Case of Low Underlying Risk

Solid blue lines in figure 3 are impulse responses to the baseline risk shock along the zero-shock path starting from the low-risk steady state. The risk shock represents a scenario in which risk is temporarily elevated so that agents expect shocks to be drawn from distributions with higher spreads for some time in the future. But the economy is not actually hit by any level shocks along this adjustment path; it is only the perception of risk that changes. When risk spikes up, agents begin to worry about the monetary policymaker's inability to respond to large adverse shocks as a consequence of

Figure 3. Impulse Responses to Risk Shocks



Notes: Impulse responses to a baseline risk shock (solid blue lines), a shock to r^* risk only (dashed red lines), and to cost-push risk only (dashed-dotted black lines), around a low-risk steady state ($\sigma = 0.0012$) in the canonical New Keynesian model with a ZLB on interest rates under optimal discretionary monetary policy.

the ZLB. Therefore, inflation expectations fall short of the inflation target, and output expectations of potential. By (1), the risk shock has a negative cost-push effect: for any given level of the output gap, inflation falls in response to lower inflation expectations. This effect induces a tradeoff for the policymaker as reflected in the targeting rule in (8). Under optimal discretion, the policymaker loosens policy enough to bring output above its efficient potential. The expansion in the economy works to limit the fall in inflation and appropriately balance deviations from target with real economic outcomes. As risk falls back, the ZLB becomes less of a concern and the economy gradually returns to the low-risk steady state.

The dynamics induced by the risk shock are similar to those following a level cost-push shock in the New Keynesian model. But with a risk shock, the interest rate has to be reduced more to achieve the optimal balance between inflation and the output gap. There are two reasons for this. First, lower inflation expectations raise the real interest rate for a given level of the nominal rate. And second, since output expectations have also been adversely affected by the risk shock, policy needs to bring about a lower real interest rate to boost aggregate demand through (2). In this sense, the increase in risk has made monetary policy less effective.

Importantly, a tradeoff arises in uncertain times even if shocks do not actually happen. The only prerequisite is that the risk shock is large enough that the ZLB becomes a concern. Small increases and reductions in risk around the low-risk steady state leave economic outcomes unaffected. Of course, the closer the economy operates to the ZLB, the more risk shocks become "large" in this sense. Reversely, if underlying risk is high, the tradeoff for monetary policy becomes a permanent feature of the economy as in the new normal described above.

7.2 On the Sources of Risk

Figure 3 also shows the effects of a positive risk shock around a low-risk steady state for each of the two shocks in turn. Again, the standard deviations for the two shock processes move together in the baseline risk shock mainly for convenience. There is no reason to rule out a priori that risk cannot move independently for the two types of shocks. Qualitatively, however, the economy responds in the same way to the two individual shocks. Spikes in risk lead to cost-push effects both when risk is elevated for r^* only (dashed red lines) and for the cost-push process only (dashed-dotted black lines). It is simply the numerical increases in risk required to induce similar quantitative dynamics that are different (top-left panel). In both cases, responses are driven by an increase in the likelihood that policymakers cannot provide sufficient stimulus. But the sources of the potential adverse shocks are immaterial. For cost-push shocks, a negative bias in inflation expectations occurs because monetary policy cannot always engineer a sufficient boom in the economy to prevent inflation from falling too much after large negative shocks. In the

case of r^* shocks, sufficiently negative realizations make it impossible for monetary policy to provide enough support for aggregate demand to keep up with supply.[16] A tradeoff arises for monetary policy as the prospect of such demand-driven recessions feed into inflation expectations when risk is elevated.[17]

Notice, however, that shocks to r_t^* are not necessarily demand shocks in the traditional sense. In the canonical New Keynesian model, fluctuations in the efficient equilibrium real rate of interest are driven by changes in the expected growth rate of total factor productivity in addition to changes in preferences and exogenous spending; see, e.g., Galí (2008). Heightened uncertainty about the future growth potential of the economy is therefore an example of a risk shock to r_t^*. A scenario in which such an increase in perceived risk is associated with a fall in expected future growth rates would correspond to a combination of a positive risk shock and a negative level shock to r^* in this framework.

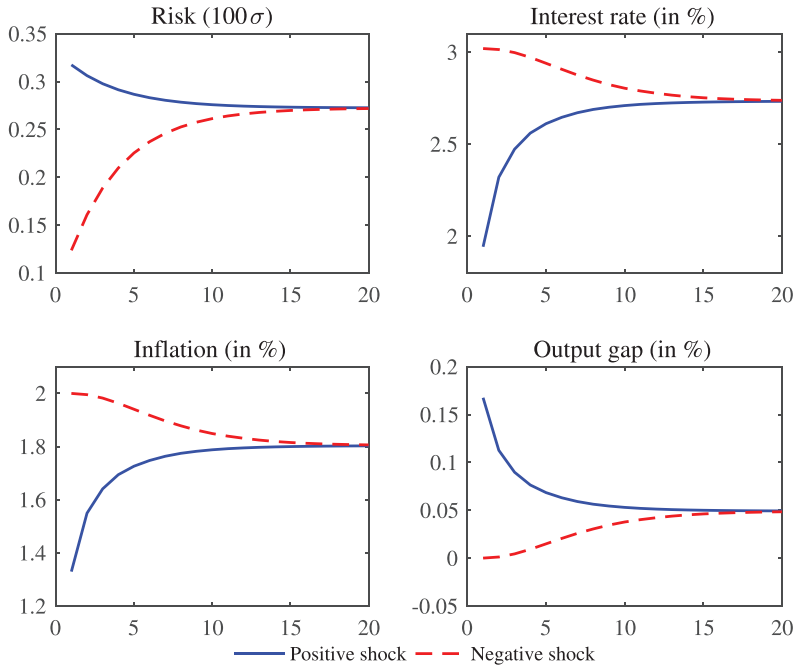
7.3 Risk Shocks in the New Normal

Starting from the new normal steady state, both positive and negative risk shocks have cost-push effects as shown in figure 4. Responses to a positive shock (solid blue lines) are as before, except that the economy reverts to the risky steady state with a negative bias in inflation. An increase in risk increases the bias in expectations, worsening the tradeoff for monetary policy. But a negative risk shock (dashed red lines) now has a positive cost-push effect. As risk falls, agents stop worrying about the ZLB, and inflation expectations realign with the inflation target. Policymakers increase interest rates in response, while the output gap closes. Gradually, as risk returns to its underlying level, the economy reverts to the high-risk steady state. As the responses show, optimal monetary policy in the

[16] If monetary policy were unrestricted by the ZLB, shocks to r_t^* could always be perfectly offset by an appropriate stance of policy. In this case, the output gap would remain closed, and inflation would be on target by the divine coincidence (Blanchard and Galí 2007).

[17] As illustrated by Adam and Billi (2007), a tradeoff arises for persistent negative *level* shocks to r^* of an intermediate size for a similar reason: when the economy moves closer to the ZLB, more future shocks can potentially cause a recession for a given level of risk.

Figure 4. Impulse Responses to Risk Shocks in New Normal

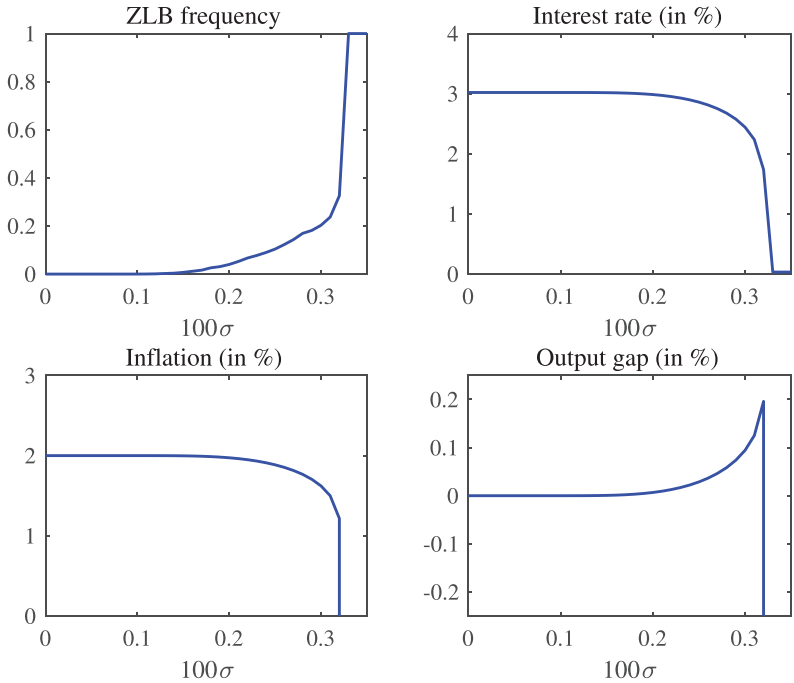


Notes: Impulse responses to a positive (solid blue lines) and a negative (dashed red lines) baseline risk shock around the new normal risky steady state ($\sigma = 0.0027$) in the canonical New Keynesian model with a ZLB on interest rates under optimal discretionary monetary policy.

new normal responds nimbly when risk perceptions change in both directions.

The asymmetry in the responses to positive and negative risk shocks around the risky steady state reflect a nonlinearity in the effect of risk on economic outcomes as illustrated in figure 5. With low levels of risk, the economy operates in the deterministic steady state in the absence of level shocks. As risk increases, the ZLB eventually becomes binding in some states of the word. For small increases, the effects are small. But as risk increases further, the frequency of ZLB episodes increases and effects begin to accelerate. Beyond a certain critical point (around $\sigma = 0.0032$ under

Figure 5. Economic Outcomes as a Function of Risk

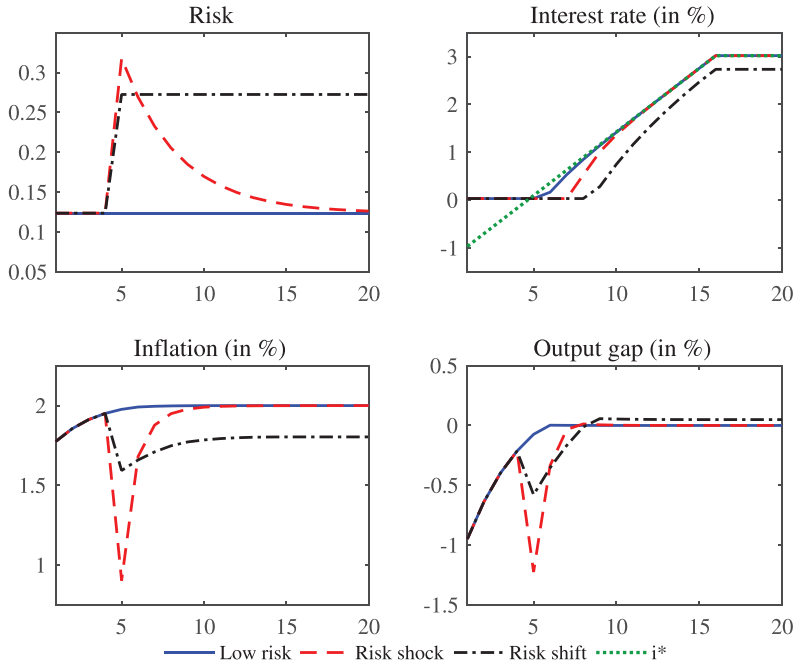


Note: Economic outcomes as a function of risk in the canonical New Keynesian model with a ZLB on interest rate under optimal discretionary policy.

the baseline calibration), interest rates are driven to the ZLB by the bias in expectations itself. In this unpleasant scenario, negative expectations—caused by a concern about the policymaker's inability to respond to adverse shocks—become self-fulfilling as the policymaker is, in fact, unable to respond sufficiently to these expectations because of the ZLB. As a result, the economy enters a downward spiral with hyperdeflation and a collapse of output.[18]

[18]The possibility of explosive dynamics corresponds to the potential nonexistence of equilibriums analyzzed by Mendes (2011) and Nakata and Schmidt (2014).

Figure 6. Normalization Scenarios



Note: Recovery from a ZLB episode as reflected in the path of the equilibrium interest rate (dotted green lines in the top-right panel) under optimal discretionary policy in a low-risk scenario (solid blue lines), in a low-risk scenario with a baseline risk shock (dashed red lines), and with a shift in risk (dashed-dotted black lines).

8. Normalization Scenarios

To illustrate the implications of a binding ZLB for the propagation of risk shocks, figure 6 shows a normalization scenario in which the economy is gradually recovering from a ZLB episode caused some time in the past by a large and persistent negative shock to the level of the equilibrium real interest rate. The nature of this initial shock—say, a financial crisis—is well understood by agents in the economy by now. Specifically, the deterministic component driven by ρ_t is known to follow the path shown in the top-right panel of figure 6 (dotted green line) so that the equilibrium nominal interest rate gradually returns to a new normal level of 3 percent. Uncertainty surrounding this recovery is perceived to be low ($\sigma = 0.0012$).

At around period $t = 4$, the efficient nominal interest rate turns positive and the policymaker, who operates under optimal discretion, is preparing to lift interest rates off the ZLB. In the absence of risk, the policymaker would simply follow the equilibrium interest rate on its trajectory back toward normal levels once it exceeds the ZLB.[19] But as long as the equilibrium interest rate is this close to the ZLB, even small shocks are "large," and the possibility that a shock may drive the economy back to the ZLB in the future is sufficient to optimally delay liftoff even when risk is low.[20]

Now suppose that agents suddenly become more uncertain about economic prospects, perhaps reflected in turmoil across financial markets. Specifically, suppose the economy is hit by a baseline risk shock corresponding to the one shown in figure 3 at time $t = 5$, just as liftoff was supposed to take place in the absence of any disturbances to the economy. Now that the economy is close to the ZLB, the impact effect of the risk shock on expectations is larger than before, as the monetary policymaker is constrained by the ZLB in its response to the shock. As shown in figure 6 (dashed red lines), inflation falls more as a consequence, and liftoff from the ZLB is further delayed. Now because of the binding ZLB, output also falls further below potential. Only as risk abates will the optimal interest rate path catch up with the equilibrium rate. The longer risk stays elevated, i.e., the more persistent the risk shock, the longer liftoff is optimally delayed even if the economy is not actually exposed to any shocks during the recovery.

Following this temporary temporary risk shock, the economy eventually returns to a low-risk steady state with inflation on target. If the shock instead takes the form of a permanent increase in underlying risk to the level associated with the new normal, the economy instead gradually settles in the risky steady state as shown in dashed dotted black lines. In this normalization scenario, optimal

[19]This corresponds to the perfect foresight case analyzed by Adam and Billi (2007) and Guerrieri and Iacoviello (2015).

[20]This is the argument made in Evans et al. (2015). But in figure 6 the ZLB binds because of an initial level shock to the equilibrium real rate of interest and not, as in their analysis, because of an explosively high risk level that may keep the economy at the ZLB for an arbitrary length of time depending on the expectational horizon.

policy lifts off from the ZLB late and continues to lean against low inflation expectations. The optimal tradeoff, however, requires the policymaker to accept that that inflation settles below below target as the economy recovers to its new normal.

9. Stochastic Volatility

So far, I have maintained the assumption that agents form expectations at any given point in time in the belief that current risk levels will persist. This assumption has allowed me to illustrate how risk interacts with the ZLB in the simplest possible framework. I now show how the results generalize to a setting in which agents understand the stochastic nature of the risk shock process.

The generalization comes at the cost of some computational complexity. As for the level shocks, I approximate the risk shock process in (6) by an independent Markov process. I assume that risk shocks are drawn first in each period, followed by the level shocks given the realization of risk. This assumption allows me to calculate one-period-ahead expectations across a three-dimensional state grid and solve the model using a generalization of the iterative procedure outlined in section 4. Specifically, I fix the state space for level shocks and calculate transition probabilities for a each level of risk using the approach in Tauchen (1986).[21] In each iteration of the solution procedure, I can find state-contingent one-period-ahead expectations conditional on the level of risk as before. In an additional step, I can now find the unconditional one-period-ahead expectations for each node in the grid as a sum of conditional expectations across risk levels weighted by transition probabilities for risk. Appendix E presents this extension to the solution method in Evans et al. (2015) in more detail.

Table 3 shows steady-state outcomes when the standard deviation of the innovation to the risk shocks process, $\nu_{\sigma,t}$, is set to

[21] I am grateful to an anonymous referee for suggesting this approach. In what follows, the Tauchen multiple parameter is set to $m = 3$ and the size of the grid for the risk shock process is $n_\sigma = 9$.

Table 3. New Keynesian Model with ZLB and Stochastic Volatility

	Interest Rate			Inflation			Output Gap			
Scenario	i^*	i	$E(i)$	π^*	π	$E(\pi)$	x^*	x	$E(x)$	P_{ZLB}
$100\sigma = 0.2668$	3.02	2.70	2.79	2.00	1.78	1.77	0.00	0.06	−0.01	0.14
$100\sigma = 0.2725$	3.02	2.65	2.76	2.00	1.75	1.74	0.00	0.06	−0.01	0.15

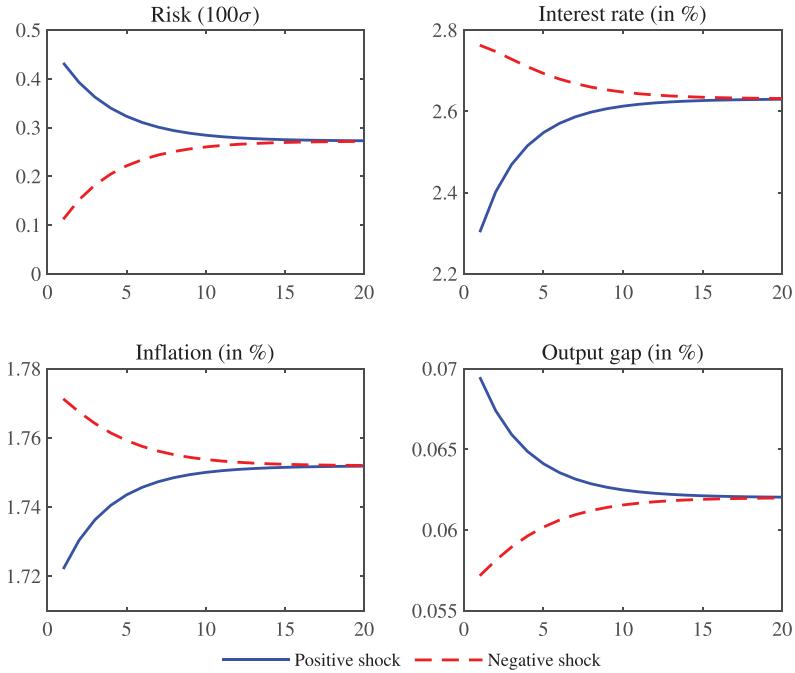
Notes: i, π, and x are risky steady-state values, stars denote deterministic steady-state values, $E(.)$ and $\sigma(.)$ denote means and standard deviations, respectively, and P_{ZLB} denotes the frequency of a binding ZLB. Deterministic steady-state values for interest rates and inflation satisfy $1 + i^*\% = (1 + r^*\%)(1 + \pi^*\%)$.

0.05/100.[22] In the first row, the underlying level of risk, $\bar{\sigma} = 0.2668/100$, is chosen so that the unconstrained model with stochastic volatility also matches the volatility of the policy rate in the hypothetical new normal discussed in section 5. Results are similar to those for the new normal considered in section 6. But despite a lower level of $\bar{\sigma}$, the bias in expectations is slightly larger and inflation settles about 2 basis points further below target in the stochastic steady state when the ZLB is imposed. In the second row, $\bar{\sigma} = 0.2725/100$ as in the new normal calibration of the model without stochastic volatility. As agents now take account of the stochastic nature of risk, inflation settles at $\pi = 1.75\%$ in the stochastic steady state, a further 5 basis points below target compared with the case with constant risk. For a given level of underlying risk, agents now find it more likely that monetary policy may be constrained in the future, as occasional spikes in risk are expected to result in larger adverse level shocks.

Figure 7 shows generalized impulse responses to a positive and a negative baseline risk shock with stochastic volatility when $\bar{\sigma} = 0.2725/100$. By contrast to the simple impulse responses shown in section 7, these impulse responses are derived under the assumption that agents expect the level of risk to follow the profile shown in the upper-left panel. Agents now have fully rational expectations about all shocks in the model. See appendix E for a precise definition. Qualitatively, the generalized impulse responses are as the same as the simple ones presented in section 7. A positive risk shock has a negative cost-push effect, and the policymaker loosens policy to limit a fall in inflation. A negative risk shock has a positive cost-push effect, and the policymaker tightens policy to limit an increase in inflation. As before, the effects of the risk shocks are asymmetric. But now that agents expect any risk shock to be temporary, the effects of a shock of a given size are much smaller. When agents expect risk to return to its underlying level according to the risk shock process, monetary policy is only expected to be constrained more often within a relatively short horizon. The effect on expectations is therefore significantly smaller. With fully rational expectations, therefore,

[22] This standard deviation ensures that risk is positive in all states for the levels of underlying risk considered.

Figure 7. Generalized Impulse Responses to Risk Shocks



Note: Impulse responses to a positive (solid blue lines) and a negative (dashed red lines) baseline risk shock around a risky steady state ($\sigma = 0.0027$) in the canonical New Keynesian model with a ZLB on interest rates and stochastic volatility under optimal discretionary monetary policy.

temporary risk shocks have to be larger to generate significant trade-offs for monetary policy.

10. Conclusion

In the canonical New Keynesian model, expectations are negatively skewed when risk is high relative to the available monetary policy space. Inflation settles materially below target in the absence of disturbances under optimal discretionary policy. Changes in the perception of risk give rise to cost-push effects regardless of the source of risk. The model is too simple to assign any great significance to quantitative results, and dynamic responses are likely to

be too immediate in the purely forward-looking framework. But the results are indicative of the direction of the effects of risk in actual economies operating in an environment in which agents have reason to worry that monetary policy may be constrained in the foreseeable future. The new normal may be one in which monetary policy should lean against a re-anchoring of inflation expectations below target by operating the economy above potential in normal times. The analysis further points toward a monetary policy strategy aiming for an inflation rate *above*, but close to, 2 percent. The results also suggest that monetary policymakers should respond nimbly to changes in the perception of risk even as the economy escapes the ZLB.

Appendix A. Optimality Conditions

The policymaker minimizes (7) subject to (1), (2), and (3), taking expectations as given so that $E_t x_{t+1} = \bar{x}^e_{t,t+1}$ and $E_t \pi_{t+1} = \bar{\pi}^e_{t,t+1}$. To solve the optimal monetary policy problem, form the Lagrangian

$$\begin{aligned}\mathcal{L} = \frac{1}{2}\left(\pi_t^2 + \kappa x_t^2\right) + \mu_{\pi,t}\left[\pi_t - \beta\bar{\pi}^e_{t,t+1} - \kappa x_t - u_t\right] \\ + \mu_{x,t}\left[x_t - \bar{x}^e_{t,t+1} + \frac{1}{\varsigma}\left(i_t - E_t\pi_{t+1} - r_t^*\right)\right] - \mu_{i,t}\left(i_t + i^*\right),\end{aligned}$$

where μ_π, μ_x, and μ_i are the multipliers on the three constraints. In addition to (1), (2), and (3), the Kuhn-Tucker conditions are

$$\begin{aligned}\pi_t + \mu_{\pi,t} &= 0 \\ \lambda x_t - \mu_{\pi,t}\kappa + \mu_{x,t} &= 0 \\ \mu_{x,t}\frac{1}{\varsigma} + \mu_{i,t} &= 0 \\ \left(i_t + i^*\right)\mu_{i,t} &= 0 \\ \mu_{i,t} &\geq 0.\end{aligned}$$

By substitution, these conditions can be reduced to

$$(i_t + i^*)(\lambda x_t + \kappa \pi_t) = 0 \tag{A.1}$$

$$\lambda x_t + \kappa \pi_t \leq 0. \tag{A.2}$$

There are two cases for condition (A.1) to consider.

Case 1. *If $\lambda x_t + \kappa \pi_t = 0$, (A.2) also holds, and the solution is determined by (1), (2), and (8). The Kuhn-Tucker conditions are all satisfied in this case if and only if $i_t + i^* \geq 0$. If it happens that $i_t + i^* = 0$, the ZLB is just binding.*

Case 2. *If $i_t + i^* = 0$, condition (3) holds, and dynamics are determined by (1) and (2) with $i_t = -i^*$. Now, we must have $\lambda x_t + \kappa \pi_t \leq 0$ for the Kuhn-Tucker conditions to be satisfied. If it happens that $\lambda x_t + \kappa \pi_t = 0$, the ZLB is just binding.*

If there exists an $i_t = i_1 > -i^*$ such that (1), (2), and (8) holds for some values $x_t = x_1$ and $\pi_t = \pi_1$, then setting $i_t + i^* = 0$ would imply that $x_t > x_1$ and $\pi_t > \pi_1$ by (1) and (2) so that $\kappa \pi_t + \lambda x_t > 0$. Hence, if the Kuhn-Tucker conditions are satisfied in case 1 given realizations of $\bar{x}^e_{t,t+1}$, $\bar{\pi}^e_{t,t+1}$, u_t, and r^*_t, case 2 will not be a candidate for a solution.

In sum, the policymaker chooses the unconstrained optimal policy allocation characterized by (1), (2), and (8) whenever it is feasible, and sets $i_t + i^* = 0$ when it is not. In the latter case, monetary policy will be insufficiently stimulatory in the sense that (A.2) holds with inequality rather than equality as in (8).

Appendix B. Solution

In each state (ϵ, u) in the $n_\epsilon \times n_u$ state space in period t, expectations are taken as given so that $E_t x_{t+1} = \bar{x}^e_{t,t+1}(\epsilon, u)$ and $E_t \pi_{t+1} = \bar{\pi}^e_{t,t+1}(\epsilon, u)$. Combining (1) and (8) in the form

$$\pi_t(\epsilon, u) = \beta \bar{\pi}^e_{t,t+1}(\epsilon, u) + \kappa x_t(\epsilon, u) + u_t(\epsilon, u)$$

$$\pi_t(\epsilon, u) = -\frac{\lambda}{\kappa} x_t(\epsilon, u)$$

gives the unconstrained optimal allocation

$$\pi_t^{opt}(\epsilon, u) = \frac{\lambda}{\lambda + \kappa^2}\left[\beta\bar{\pi}_{t,t+1}^e(\epsilon, u) + u_t(\epsilon, u)\right]$$
$$x_t^{opt}(\epsilon, u) = -\frac{\kappa}{\lambda + \kappa^2}\left[\beta\bar{\pi}_{t,t+1}^e(\epsilon, u) + u_t(\epsilon, u)\right].$$

The interest rate consistent with this allocation follows from (2):

$$i_t^{opt}(\epsilon, u) = \bar{\pi}_{t,t+1}^e(\epsilon, u) + r_t^*(\epsilon, u) - \sigma\left[x_t^{opt}(\epsilon, u) - \bar{x}_{t,t+1}^e(\epsilon, u)\right].$$

If $i_t^{opt}(\epsilon, u) \geq -i^*$, $\left\{x_t^{opt}(\epsilon, u), \pi_t^{opt}(\epsilon, u), i_t^{opt}(\epsilon, u)\right\}$ is the solution for state (ϵ, u) in period t. If the ZLB is binding so that $i_t^{opt}(\epsilon, u) < -i^*$, the interest rate is set to $i_t^{zlb}(\epsilon, u) = -i^*$. Now from (2) and (1):

$$x_t^{zlb}(\epsilon, u) = \bar{x}_{t,t+1}^e(\epsilon, u) - \frac{1}{\sigma}\left[-i^* - \bar{\pi}_{t,t+1}^e(\epsilon, u) - r_t^*(\epsilon, u)\right]$$
$$\pi_t^{zlb}(\epsilon, u) = \beta\bar{\pi}_{t,t+1}^e(\epsilon, u) + \kappa x_t^{zlb}(\epsilon, u) + u_t(\epsilon, u).$$

Hence, the solution for $y_t(\epsilon, u) \in \{x_t(\epsilon, u), \pi_t(\epsilon, u), i_t(\epsilon, u)\}$ for all nodes (ϵ, u) in the state grid is

$$y_t^{sol}(\epsilon, u) = \begin{cases} y_t^{opt}(\epsilon, u) & \text{if } i_t^{opt}(\epsilon, u) \geq -i^* \\ y_t^{zlb}(\epsilon, u) & \text{if } i_t^{opt}(\epsilon, u) < -i^*. \end{cases}$$

Ex ante expectations across the state grid can now be found as

$$\bar{\mathbf{x}}_{t-1,t}^e = \mathbf{P}_\epsilon \mathbf{x}_t^{sol} \mathbf{P}_u'$$
$$\bar{\pi}_{t-1,t}^e = \mathbf{P}_\epsilon \pi_t^{sol} \mathbf{P}_u',$$

where $\mathbf{P}_\epsilon$ and $\mathbf{P}_u$ are Markov transition matrices of dimensions $n_\epsilon \times n_\epsilon$ and $n_u \times n_u$, respectively.

The solution algorithm iterates the solution backwards from some period $t = T \gg 0$ to $t = 0$, initiated with $\bar{x}_{T,T+1}^e = \bar{\pi}_{T,T+1}^e = \mathbf{0}$. The state-contingent model solution is then $\Big\{x_0^{sol}(\epsilon, u), \pi_0^{sol}(\epsilon, u), i_0^{sol}(\epsilon, u)\Big\}$.

Appendix C. Calculation of Model Statistics

The stochastic steady state of model variable $y_t \in \{i_t, \pi_t, x_t\}$ is simply

$$y \equiv y_0^{sol}(\epsilon = 0, u = 0).$$

To calculate unconditional expectations, note that for each shock process $z_t \in \{\epsilon_t, u_t\}$ with Markov transition matrix $\mathbf{P}_z$, the stationary unconditional distribution, $\mathbf{d}_z$, satisfies $\mathbf{d}_z' = \mathbf{d}_z'\mathbf{P}_z$, or equivalently $(\mathbf{I} - \mathbf{P}_z')\mathbf{d}_z = 0$. Hence, the unconditional distribution can be found as the normalized eigenvector associated with the unitary eigenvalue of $\mathbf{P}_z'$. The unconditional probability distribution over the state space is then the $n_\epsilon \times n_u$ vector product $\mathbf{D} = \mathbf{d}_\epsilon \mathbf{d}_u'$. The unconditional expectation of model variable y_t is found as

$$E(y) \equiv \sum_\epsilon \sum_u D(\epsilon, u) y_0^{sol}(\epsilon, u).$$

The unconditional variance is

$$\sigma^2(y) \equiv \sum_\epsilon \sum_u D(\epsilon, u) \left[y_0(\epsilon, u) - E(y)\right]^2.$$

Appendix D. Solution with Stochastic Volatility

With stochastic volatility, the discrete state space is extended to the $n_\epsilon \times n_u \times n_\sigma$ grid (ϵ, u, σ). Each period, the risk shock is assumed to be drawn first, followed by the level shocks given the realization of risk. With this timing assumption, a conventional method can be used to approximate the risk shock process (6) for a given volatility of $\nu_{\sigma,t}$, resulting in a $n_\sigma \times 1$ state space and an associated Markov transition matrix $\mathbf{P}_\sigma$. Let the set of grid points be $\tilde{\sigma} = \{\tilde{\sigma}_1, \tilde{\sigma}_2, \ldots, \tilde{\sigma}_{n_\sigma}\}$, centered at $\tilde{\sigma}_i = \bar{\sigma}$ for $i = n_\sigma - (n_\sigma - 1)/2$. The $n_z \times 1$ state space for each $z_t \in \{\epsilon_t, u_t\}$ can then be selected following the approach in Tauchen (1986) for the underlying level of risk, $\bar{\sigma}$. Assuming $\varsigma = 1$, the maximum value in the grid is set to

$$z_{n_z} = m \left(\frac{\bar{\sigma}^2}{1 - \mu_z} \right)^{\frac{1}{2}}$$

for some $m \in Z^+$, the minimum value to $z_1 = -z_{n_z}$, and the distance between states to $w = (z_{n_z} - z_1)/(n_z - 1)$. While the state space itself is kept fixed, transition probabilities now depend on the realization of risk. Each element in each Markov transition matrix $\mathbf{P}_z^\sigma = (p_{z,jk}^\sigma)_{n_z \times n_z}$ for each state of σ_t can be found using Tauchen's (1986) formula

$$p_{z,jk}^\sigma = \begin{cases} F\left(\frac{z_k - \mu_z z_j + w/2}{\sigma}\right) & \text{if } k = 1 \\ F\left(\frac{z_k - \mu_z z_j + w/2}{\sigma}\right) & \\ \quad -F\left(\frac{z_k - \mu_z z_j - w/2}{\sigma}\right) & \text{if } k \in \{2, 3, \ldots, n_z - 1\} \\ 1 - F\left(\frac{z_{n_z} - \mu_z z_j - w/2}{\sigma}\right) & \text{if } k = n_z, \end{cases}$$

where $F(.)$ is the standard normal cumulative probability distribution.

Expectations are taken as given in each state in period t so that $E_t y_{t+1} = \bar{y}_{t,t+1}^e(\epsilon, u, \sigma)$. The state-contingent solution $y_t^{sol}(\epsilon, u, \sigma)$ conditional on these expectations as well as the ex ante expectations conditional on the level of risk, $\bar{\mathbf{y}}_{t-1,t}^{e|\sigma}$, can be found as outlined in appendix B. But the calculation of the unconditional ex ante expectations, $\bar{y}_{t-1,t}^e(\epsilon, u, \sigma)$, requires an additional step with stochastic volatility. With a slight abuse of notation, these expectations can be found as

$$\bar{y}_{t-1,t}^e(\epsilon, u, \sigma = \tilde{\sigma}_i) = \sum_j p_{\sigma,ij}\, \bar{y}_{t-1,t}^{e|\sigma_j}(\epsilon, u),$$

where $p_{\sigma,ij}$ denotes the ij-th element of $\mathbf{P}_\sigma$.

The state-contingent model solution with stochastic volatility can now be found by iterating the period-t solution, $y_t^{sol}(\epsilon, u, \sigma)$, backwards from some period $t = T \gg 0$ to $t = 0$, initiated with $\bar{x}_{T,T+1}^e(\epsilon, u, \sigma) = \bar{\pi}_{T,T+1}^e(\epsilon, u, \sigma) = 0$, and calculating $\bar{x}_{t-1,t}^e(\epsilon, u, \sigma)$ and $\bar{\pi}_{t-1,t}^e(\epsilon, u, \sigma)$ in subsequent iterations, until convergence.

Model statistics for the case of stochastic volatility are calculated using straightforward extensions of the formulas in appendix C to three dimensions with $D(\epsilon, u, \sigma) = d_\sigma(\sigma) D(\epsilon, u \mid \sigma)$.

Appendix E. Risk Shocks with Stochastic Volatility

Following Koop, Pesaran, and Potter (1996), the generalized impulse response function for the risk shock process is defined as

$$\begin{aligned} GI^{y}_{t+n} = {} & E_t(y_{t+n} \mid \sigma_t = \tilde{\sigma}_i, \epsilon_t = 0, \ldots, \epsilon_{t+n} = 0, \\ & \quad u_t = 0, \ldots, u_{t+n} = 0) \\ & - E_t(y_{t+n} \mid \sigma_t = \bar{\sigma}, \epsilon_t = 0, \ldots, \epsilon_{t+n} = 0, u_t = 0, \ldots, u_{t+n} = 0) \end{aligned}$$

when $\tilde{\sigma}_i \in \tilde{\sigma}$ is the impact value of risk following the period-t innovation. This definition maintains the assumption that risk never materializes in nonzero-level disturbances along the adjustment path. By the Markov chain approximation we have

$$\begin{aligned} GI^{y}_{t+n} = {} & \sum_j p^n_{\sigma,ij}\, y^{sol}_0(\epsilon = 0, u = 0, \sigma = \tilde{\sigma}_j) \\ & - \sum_j p^n_{\sigma,mj}\, y^{sol}_0(\epsilon = 0, u = 0, \sigma = \tilde{\sigma}_j), \end{aligned}$$

where $m = n_\sigma - (n_\sigma - 1)/2$ and $p^n_{\sigma,ij}$ denotes the ij-th element of $\mathbf{P}^n_\sigma$ (i.e., the n-th power of the transition matrix $\mathbf{P}_\sigma$).

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Independent Central Banks: Low Inflation at No Cost? A Model with Fiscal Policy*

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In this article we extend the rational partisan model of Alesina and Gatti (1995) to include a second policy, fiscal policy, besides monetary policy. It is shown that the extent to which an independent central bank is successful in attaining price stability depends on the degree of conservativeness of the central bank in relation to the political parties and the private sector's expectations on which party will win the elections. In addition, the inclusion of fiscal policy in Alesina and Gatti's model implies that uncertainty about the course of policy is not a sufficient factor to ensure that, when supply shocks are not relevant, independent central banks bring about low inflation at no real cost.

JEL Codes: E58, E63.

1. Introduction

By appointing a conservative independent central bank to take control of monetary policy, Rogoff (1985) showed that average inflation would be reduced. Given the tradeoff between the objectives of output and inflation stabilization, a conservative central bank would prioritize fighting inflation with a theoretical cost of higher output variability. Alesina and Gatti (1995), based on the lack of empirical evidence of higher output variability shown by Alesina

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and Summers (1993), developed a theoretical model to illustrate why a conservative independent central bank might not bring higher volatility of output.

The rational partisan model presented by Alesina and Gatti (1995) included two political parties running for office, with different views of the economy. Monetary policy, the only policy in their model, is either decided by the party that wins the elections or delegated to an independent central bank. The authors identify two sources of instability: the first source is due to the uncertainty about which party will be in office and the second one is due to exogenous shocks. Consequently, they decompose the variability of output into two components: the political volatility, introduced by the uncertainty about the future course of policy, and the economic volatility, induced by exogenous shocks. Alesina and Gatti (1995) show that by removing the conduct of monetary policy from the hands of the government, the first component of the variance of output is eliminated, allowing for the possibility that the overall volatility of output does not necessarily increase. In particular, these authors conclude that if the volatility of shocks is low enough, delegation of the conduct of monetary policy reduces the overall variance of output.

Monetary policy has been considered an ideal candidate for delegation (see, for instance, Drazen 2002 and Alesina and Tabellini 2007, 2008), due to its technical nature and the difficulty in judging the ability or talent of the person responsible for making the decisions. Fiscal policy, on the other hand, is not viewed as a clear candidate for delegation, mainly because of its redistributive impact. Moreover, as fiscal policy can secure a minimum number of voters, politicians will not willingly delegate such policy if they want to be re-elected. Therefore, fiscal and monetary policies are implemented in many countries by different authorities that are generally independent from each other. For this reason, an interesting extension of Alesina and Gatti (1995) would be the inclusion of fiscal policy in the model, in order to see whether an independent central bank responsible for monetary policy and presumably isolated from electoral cycles is still able to eliminate the politically induced volatility of output.

In this paper, we generalize Alesina and Gatti's model by introducing a second policy, fiscal policy, that will be decided by the party in government. We initially consider a basic framework where

(i) the two parties running for office only differ in the relative weights assigned to output stabilization and (ii) the government and the central bank simultaneously choose their policy in case of delegation of monetary policy. In the next paragraphs we explain the main results obtained.

The benefits in terms of inflation (low and stable inflation) of the appointment of an independent central bank depend on the degree of conservativeness of the central bank in relation to the political parties and the private sector's expectations on which party will win. An ultraconservative independent central bank (i.e., a central bank more conservative than the two parties) is always expected to achieve lower and more stable inflation. However, a moderately independent central bank (i.e., a central bank that has an inflation aversion intermediate between the two parties) is expected to attain lower and more stable inflation only if the probability of the less inflation-averse party winning the elections is high enough.

Further, the politically induced output variability is not removed by the introduction of the central bank—in fact, it may even increase.[1] This last case occurs when the central bank is ultraconservative and political parties are not particularly concerned about achieving their public spending targets. In this case, when monetary policy is delegated to an independent ultraconservative central bank, the inflation rate is chosen almost regardless of output deviation. This will lead to a higher politically induced volatility of output compared with nondelegation, where outputs will be closer to the target and will differ less.

Similarly to Alesina and Gatti's results, the appointment of an ultraconservative central bank unequivocally increases the economically induced variance of output. By contrast, when a moderately conservative independent central bank is responsible for monetary

[1]In Alesina and Gatti's model, as there is no fiscal policy, the central bank chooses the same inflation under delegation, regardless of the party that is in office. This, in turn, leads to the same value of output and, therefore, one concludes that the institution of an independent central bank eliminates the politically induced output variability in their model. By contrast, when fiscal policy is introduced, the central bank will choose the inflation level depending on the party that is in office. This implies that the level of output obtained depends on the party that is in office and, therefore, the politically induced output variability is present in our model.

policy, this component of the variance of output increases whenever the probability of the less inflation-averse party winning the elections is high enough.

In addition, the relationship between central bank independence and output variability depends on the degree of conservativeness of the central bank and on exogenous shocks. We will show the conditions that will lead to an increase or a decrease of the overall variance of output. In this way, our analysis suggests that if the volatility of shocks is low enough, delegation of the conduct of monetary policy may not reduce the overall variance of output and, hence, the above-mentioned conclusion reached by Alesina and Gatti is not robust.

The last result we would like to point out is that, in contrast to Alesina and Gatti (1995), the study of output stabilization in our model is not reduced to the study of the variance of output. Unlike their paper, we obtain that output is not necessarily more stable under delegation of monetary policy to an independent central bank when economic shocks are not relevant. For instance, when the parties' preferences are not too different, output is less stable with an ultraconservative independent central bank.

To test the robustness of our results, we analyze two generalizations of the basic framework: (i) parties differ in their target for public spending, and (ii) the authorities choose their policies sequentially. In relation to the first extension of the model, recent empirical studies have provided evidence to support the popular view that left-wing parties are associated with higher public spending (see, for instance, Blomberg and Hess 2003, Pickering and Rockey 2011, 2013, among others). Consequently, we generalize the model, assuming that a left-wing party would prefer a higher target for government expenditure than a right-wing party would. In relation to the second generalization, one could argue that as the process of changing tax rates takes longer than the process to adjust monetary policy, a more appropriate description of fiscal–monetary interactions would involve a leader–follower game. In this second variation of the basic framework, the fiscal authority acts as the leader and the monetary authority acts as the follower. We show that the main results derived in the basic framework also hold in these new setups.

The different outcomes delivered by our model can account for the mixed results obtained by the empirical literature. The

initial evidence seemed to favor the existence of a negative relationship between central bank independence (CBI) and inflation in OECD countries (see, for instance, the surveys of Eijffinger and de Haan 1996 and Berger, de Haan, and Eijffinger 2001, or the meta-regression analysis of Klomp and de Haan 2010b), but for developing countries the situation is less clear. In fact, when large heterogeneous samples of countries are used, no general significant negative relation between CBI and inflation is found (Klomp and de Haan 2010a and Dincer and Eichengreen 2014).[2] The situation is similar for the relationship between CBI and growth. While some studies fail to find a robust relationship between CBI and economic growth (Cukierman et al. 1993, Eijffinger, van Rooij, and Schaling 1996, Akhand 1998, Eijffinger, Schaling, and Hoeberichts 1998, and de Haan and Kooi 2000, for instance), others point out that CBI increases the variability of output (Fujiki 1996, Fuhrer 1997, Hall and Franzese 1998, and Zervoyianni, Anastasiou, and Anastasiou 2014).

One of the reasons suggested for the discordant empirical results is related to the variables used to measure central bank independence. The de facto independence of the central bank is not properly captured by the legal definition. In addition, the presence of a relationship between CBI and inflation or growth does not imply causality.[3] Further, the distinction between independence and the degree of conservativeness or inflation aversion presented in theoretical models is not easily captured by the variables used empirically. It will be shown in this paper that the effects of CBI on inflation and output stabilization will be dependent on the degree of central bank conservativeness.

The present paper is related to the theoretical literature that focuses on the impact of delegating monetary policy to an independent central bank. Demertzis (2004) carries out numerical simulations of Alesina and Gatti's (1995) model and shows that changing political uncertainty values could alter their results. Further, by introducing fiscal policy, this article can be associated to the

[2]Dincer and Eichengreen (2014) obtain some—although statistically inconsistent—negative relationship.

[3]More on this can be found in, among others, Cukierman et al. (1993), Alpanda and Honig (2010), Klomp and de Haan (2010b), and Dincer and Eichengreen (2014).

literature that studies the interaction of monetary and fiscal policy (see, for instance, Alesina and Tabellini 1987, Debelle and Fisher 1994, Beetsma and Bovenberg 1997, Dixit and Lambertini 2003, among others). However, our model extends this literature by introducing electoral uncertainty. In a previous article, Ferré and Manzano (2014) have included electoral uncertainty in a model with two policymakers by extending Alesina's (1987) rational partisan theory model. It is shown that the inclusion of a central bank can alter the predictions of the rational partisan theory, in the sense that the direct relationship predicted between inflation and output in Alesina (1987) does not hold.

The remainder of this paper is organized as follows. The next section will develop a rational partisan model where fiscal and monetary policies are initially under the control of the government and, then, monetary policy is delegated thereafter to an independent central bank. Additionally, the effects of the introduction of an independent central bank, responsible for monetary policy, on expected inflation and inflation stabilization are analyzed. Section 3 checks the robustness of the model by including two extensions of the basic setup. Finally, section 4 will present the conclusions.

2. The Benchmark Model

In this section, we will present a model that combines features from both Alesina and Tabellini (1987) and Alesina and Gatti (1995). We will assume that there are two parties competing for office, L (a left-wing party) and R (a right-wing party), and there is an exogenous probability P that party L wins the elections and takes office. Agents (wage setters) in this economy will not know what party will be in office when they form their inflation expectations, π^e. For this reason, their expectations embody electoral uncertainty: $\pi^e = PE(\pi_L) + (1-P)E(\pi_R)$, where $E(\pi_j)$ represents expected inflation if party j is in office ($j = L, R$). Once elections take place, the party in office will attempt to stabilize the economy after the shocks occur, and the optimal values of inflation and taxes will be revealed. This sequential structure of the game is static by nature, as the game ends once the policy instruments are chosen.

If party j is in office, the output is given by

$$x_j = \pi_j - \pi^e - \tau_j - w^* + \varepsilon, \tag{1}$$

where π_j is the actual inflation rate.[4] Moreover, τ_j represents taxes levied on output, w^* denotes the target real wage that workers seek to achieve, and ε is a productivity shock such that $E(\varepsilon) = 0$ and $var(\varepsilon) = \sigma_\varepsilon^2$.

The budget constraint of government j is

$$g_j = \tau_j + \pi_j, \tag{2}$$

where g_j denotes the ratio of public expenditures over output when party j is in office. Note that public spending will be financed by a distortionary tax (controlled by the fiscal authority) and/or by money creation (controlled by the authority responsible for monetary policy). Given the static nature of the model, debt is not included.[5]

We assume that the loss function for party j is given by

$$V_{Gj} = \frac{1}{2}\left(\pi_j^2 + \delta_j(x_j - x^*)^2 + \gamma(g_j - g^*)^2\right), \tag{3}$$

where δ_j and γ represent the relative weights assigned to output and public spending stabilization with respect to inflation, respectively, and δ_j, $\gamma > 0$, while x^* and g^* denote the output and public spending targets, respectively. The government's objective function is given by an augmented but otherwise conventional loss function (see, for instance, Alesina and Tabellini 1987, Debelle and Fisher 1994, Beetsma and Bovenberg 1997, Huang and Wei 2006, and Hefeker 2010, among others). This objective function reflects that the government aims to stabilize output and inflation simultaneously, as well as meet a spending target, which could reflect the

[4] A detailed derivation of expressions (1) and (2) is given at the beginning of appendix A.

[5] By not including debt in the model we avoid the introduction of another interesting but separate issue, namely the manipulation of economic variables to influence the outcome of elections. The incumbent party could affect the result of the elections and/or the success of the mandate of the winning party by increasing spending and debt.

aim of being re-elected or other demands from interest groups that influence the government. Following the literature, we suppose that $\delta_L > \delta_R$. In the benchmark case, we assume that parties have identical relative weights assigned to public spending stabilization and share the same goals. This framework will allow us to make a clear comparison between our results and the ones derived by Alesina and Gatti (1995).[6]

In what follows we distinguish two frameworks: first, when monetary and fiscal policy are controlled by the government, and second, when monetary policy is delegated to an independent authority (central bank). The first framework will represent an economy with no (or very little) central bank independence, whereas the second one will refer to an economy that has granted independence to its central bank for the conduct of monetary policy. In both cases, the timing of events is as follows: expectations and, thus, wages are set first. Afterward, elections take place; party L wins with probability P, and party R with probability $1 - P$. After the election, the shock ε occurs. In the first case, the government chooses both policies. In the second case, the government and the central bank will simultaneously choose their policy.

2.1 No Independent Monetary Policy

When monetary and fiscal policy are both under the control of the government, the party in government will attempt to minimize its loss function (3) by using two instruments, π and τ. The inflation rates chosen by the two parties if in office and the corresponding outputs are (where the superscript N indicates nondelegation of monetary policy)

$$\pi_L^N = \frac{m_R + 2}{\Delta^N} A - \frac{\varepsilon}{m_L + 2}, \tag{4}$$

$$\pi_R^N = \frac{m_L + 2}{\Delta^N} A - \frac{\varepsilon}{m_R + 2}, \tag{5}$$

$$x_L^N = x^* - \frac{1}{2\delta_L} \pi_L^N, \text{ and} \tag{6}$$

[6]We would like to thank an anonymous referee for this suggestion.

$$x_R^N = x^* - \frac{1}{2\delta_R}\pi_R^N, \tag{7}$$

where $m_L = \frac{\frac{1}{\delta_L}+\frac{1}{\gamma}}{2}$, $m_R = \frac{\frac{1}{\delta_R}+\frac{1}{\gamma}}{2}$, $\Delta^N = (m_L+2)(m_R+1) + P(m_L - m_R)$, and $A = x^* + g^* + w^*$.[7]

As it is indicated in Ferré and Manzano (2014), m_j represents a measure of party j's inflation aversion.[8] The assumption that $\delta_L > \delta_R$, that is, party L gives more weight to output stabilization than party R, implies that $m_R > m_L$, i.e., the goal of stabilizing inflation is more important for party R than for party L.[9] Accordingly, taking expectations in expressions (4) and (5), it follows that $E(\pi_L^N) > E(\pi_R^N)$ as $m_R > m_L$, i.e., expected inflation will always be higher under an L administration.

The following lemma shows that in the benchmark model expected output will be also higher under an L administration.

Lemma 1. $E(x_L^N) > E(x_R^N)$.

According to lemma 1, we expect a lower deviation of output when party L is in office. This is due to the fact that party L is more concerned about output stabilization than party R.

2.2 Introducing an Independent Monetary Authority

We will now study the case where monetary policy is undertaken by an independent monetary authority. Independence refers to the extent to which the central bank determines monetary policy without political interference. Hence, when party j is in office, we will assume now that the central bank will have its own loss function to minimize given by

[7] A detailed derivation of the optimal policies under nondelegation and delegation of monetary policy to an independent central bank can be found in the appendix (see propositions A.1 and A.2, respectively).

[8] Notice that 1 is the weight attributed to inflation in the parties' loss functions. Thus, m_j is the arithmethic mean of the weight of inflation relative to output and public spending for party j.

[9] In models with only one policy (and, in particular, $\gamma = 0$), it is assumed that $\delta_L > \delta_R$ —see, for instance, Alesina (1987) and Alesina and Gatti (1995)—and, thus, in this case $m_L = 1/\delta_L$ and $m_R = 1/\delta_R$, which would also correspond to $m_R > m_L$.

$$V_{CB} = \frac{1}{2}\left(\pi_j^2 + \delta_{CB}\left(x_j - x^*\right)^2\right), \tag{8}$$

where $\delta_{CB} > 0$. In this case, the timing of events is the same, but after the shock ε occurs, the central bank will control inflation (π) to minimize its loss function (8), and the party in government will attempt to minimize its loss function (3) by using taxes (τ). With this institutional specialization we obtain the following inflation rates and outputs (where superscript D indicates delegation of monetary policy):

$$\pi_L^D = \frac{c_R m_R + 2}{\Delta^D} A - \frac{\varepsilon}{c_L m_L + 2}, \tag{9}$$

$$\pi_R^D = \frac{c_L m_L + 2}{\Delta^D} A - \frac{\varepsilon}{c_R m_R + 2}, \tag{10}$$

$$x_L^D = x^* - \frac{c_L}{2\delta_L}\pi_L^D, \text{ and} \tag{11}$$

$$x_R^D = x^* - \frac{c_R}{2\delta_R}\pi_R^D, \tag{12}$$

where $\Delta^D = \left(c_L m_L + 2\right)\left(c_R m_R + 1\right) + P\left(c_L m_L - c_R m_R\right)$. In these expressions we have introduced two new variables, c_L and c_R. The variable c_j, $j = L, R$, is a measure of the degree of the relative conservativeness of the central bank with respect to party j:[10] $c_j = \frac{2\delta_j}{\delta_{CB}}$.

Remark 1. *If $c_L = 1$ and $c_R = 1$, that is, the central bank is as conservative as both parties, then $\pi_j^D = \pi_j^N$ and $x_j^D = x_j^N$.*

Using the expressions for m_L, m_R, c_L, and c_R, we have that $c_L m_L > c_R m_R$. Hence, expressions (9) and (10) imply that $E\left(\pi_R^D\right) > E\left(\pi_L^D\right)$. If party L is relatively more interested in stabilizing output than party R, party L is expected to have more incentives to reduce taxes. This, in turn, has an effect on the behavior of the central bank: the decrease in taxes diminishes the incentives to inflate and, thus, $E\left(\pi_R^D\right) > E\left(\pi_L^D\right)$.

[10]The notion of conservativeness generally refers to the degree of the central bank's inflation aversion. See Ferré and Manzano (2012) for a detailed explanation of the conservativeness measure c.

If we compare expected outputs in the presence of an independent central bank, we obtain the following result:

LEMMA 2. $E\left(x_L^D\right) > E\left(x_R^D\right)$.

Rogoff (1985) showed that, in a model with only monetary policy, society's welfare could be improved by appointing a more conservative central bank. We will follow the Rogoff tradition and assume that an agreement can be reached to appoint a central bank that is more conservative than both political parties: $c_L > c_R \geq 1$. We will refer to this central bank as being "ultraconservative." Alesina and Gatti (1995) point out that if political parties are polarized, it might not be easy to reach an agreement to delegate the conduct of monetary policy to an independent institution. They argue, however, that such an agreement will be easier to reach when the independent institution has an inflation aversion that is intermediate. Following these authors, we will also analyze a central bank more conservative than the left-wing party and less conservative than the right-wing party. In our framework this assumption is represented by $c_L > 1 > c_R$ and we label it "moderately conservative."

According to Alesina and Gatti (1995), "the institution of an independent and inflation-averse central bank has two benefits: first, it reduces average inflation; second, it eliminates politically induced output variability." In the following two subsections, we will analyze whether these results hold when the model is extended to consider two policies.

2.3 The Effects of an Independent Central Bank on Inflation

What is the effect of the introduction of an independent central bank, responsible for monetary policy, on expected inflation and inflation stabilization? The following proposition shows that it will depend on the degree of conservativeness of the central bank.

PROPOSITION 1. *(a) By appointing an ultraconservative ($c_L > c_R \geq 1$) independent central bank responsible for monetary policy, the expected value of inflation is reduced and a higher degree of inflation stabilization is achieved. (b) By appointing a moderately conservative ($c_L > 1 > c_R$) independent central bank, the expected value of inflation is reduced and a higher degree of inflation stabilization is achieved if and only if P is high enough.*

In the presence of political uncertainty, inflation is generally lower (in expected terms) and more stable when monetary policy has been delegated to an independent and conservative central bank. In other words, this proposition indicates that when monetary policy is carried out by a conservative and independent central bank, agents expect inflation to be lower and more stable than if monetary policy was set by the parties. An exception arises in case (b), where the central bank is less conservative than party R. When the probability of party R coming to power is high enough (that is, P is low enough), we expect lower and more stable inflation without delegating monetary policy to an independent central bank, as this party is already very inflation averse.[11]

2.4 The Effects of an Independent Central Bank on Output

The theoretical research that followed Rogoff's article (1985) suggested that central bank independence came at a cost of higher output variability. However, as empirical studies did not seem to find clear evidence of a higher variance of output, Alesina and Gatti (1995) developed a model where they decomposed the variance of output in two parts: the *politically induced variance* (Var_P), which reflects the fluctuations in the variable induced by electoral uncertainty, and the *economically induced variance* (Var_E), which is due to the exogenous shocks. In Alesina and Gatti (1995), removing the conduct of monetary policy from the government eliminates the politically induced variance of output, which explains why the variance of output might not necessarily increase with an independent conservative central bank. We will study how this result is altered with the introduction of fiscal policy in the analysis.

2.4.1 The Politically Induced Variance of Output

The politically induced variances of output when monetary policy is under the control of the government (N) and when it

[11]Demertzis (2004) carries out numerical simulations on Alesina and Gatti's model and also finds that, for an intermediate central bank, inflation might not always be lower.

is conducted by an independent central bank (D) are given by[12]

$$Var_P(x^N) = P\,(1-P)\left(E(x_L^N) - E(x_R^N)\right)^2 \text{ and} \tag{13}$$

$$Var_P(x^D) = P\,(1-P)\left(E(x_L^D) - E(x_R^D)\right)^2. \tag{14}$$

The last expression implies that, in general, the politically induced variance of output does not vanish when monetary policy is delegated to an independent central bank. There will only be two scenarios in which this variance vanishes: when there is no political uncertainty ($P = 0, 1$) and when $E(x_L^D) = E(x_R^D)$, which occurs when $\delta_L = \delta_R$ or $\gamma = 0$. In these last cases both parties behave identically and, consequently, the political uncertainty introduced by elections does not play any role. These results are summarized in the following proposition:

Proposition 2. *The appointment of an independent central bank when there is more than one policy instrument does not eliminate the variance of output induced by political uncertainty, except when both parties behave identically (i.e., $\delta_L = \delta_R$ or $\gamma = 0$).*

Given that the politically induced variance of output is not automatically eliminated by introducing an independent central bank, in the next lines we will study whether this variance is at least reduced with delegation of monetary policy.

Notice that the comparison of $Var_P(x^N)$ and $Var_P(x^D)$ in (13) and (14) is equivalent to contrasting the distance between the expected values of output, $\left|E(x_L^i) - E(x_R^i)\right|$, with $i = N, D$. Consequently, the comparison of this type of variances is reduced to the study of expected outputs under both frameworks N and D.

Further, nondelegation and delegation would coincide when $c_L = c_R = 1$, i.e., when monetary policy is undertaken by a central bank that is as conservative as the two parties. We can then study the effect of moving toward a moderately conservative central bank ($c_L > 1 > c_R$) by analyzing the impact of increasing the relative conservativeness of the central bank with respect to party L (an

[12]Lemma A.1 in the appendix shows the derivation of these expressions.

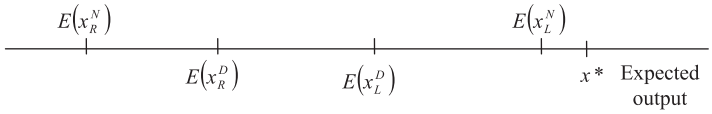
increase in c_L) and lowering the relative degree of conservativeness of the central bank with respect to party R (a decrease in c_R). Similarly, we can study the effect of introducing an ultraconservative central bank ($c_L > c_R \geq 1$) by analyzing the consequences of increasing the relative conservativeness of the central bank with respect to both parties (an increase in both c_L and c_R). The following result will prove useful in explaining how expected outputs are affected in moving from N (no independent central bank) to D (independent central bank):

Lemma 3. *Let $i, j = L, R$, $i \neq j$. A change in the relative conservativeness of the central bank with respect to party j will have two effects on expected outputs: the direct effect ($\frac{\partial}{\partial c_j} E(x_j^D)$) and the indirect effect ($\frac{\partial}{\partial c_j} E(x_i^D)$). Moreover, it holds that $\frac{\partial}{\partial c_j} E(x_j^D) < 0$ and $\frac{\partial}{\partial c_j} E(x_i^D) > 0$.*

The logic of lemma 3 is as follows. Without any loss of generality, let's start from the initial nondelegation situation ($(c_L, c_R) = (1, 1)$) and assume an increase in c_L, keeping c_R constant ($c_L > c_R = 1$). This corresponds to a new situation identical to the initial one, except that now monetary policy is undertaken by a more conservative central bank if party L is in office. As lemma 3 points out, this change in c_L will have two effects on expected outputs: a direct effect $\left(\frac{\partial}{\partial c_L} E\left(x_L^D\right)\right)$ and an indirect effect $\left(\frac{\partial}{\partial c_L} E\left(x_R^D\right)\right)$. Under the direct effect, as the authority in charge of monetary policy becomes more conservative, the difference between expected inflation and expected average inflation $(E\left(\pi_L\right) - \pi^e)$ becomes smaller, and thus expected output under party L's office will be lower. Hence, $\frac{\partial}{\partial c_L} E\left(x_L^D\right) < 0$. Under the indirect effect, the possibility that the authority in charge of monetary policy under the other party's office (party L) is more conservative will bring expected average inflation (π^e) down. Thus, the difference between expected inflation and expected average inflation $(E\left(\pi_R\right) - \pi^e)$ becomes larger and, hence, expected output under party R's office will be higher. Therefore, $\frac{\partial}{\partial c_L} E\left(x_R^D\right) > 0$.[13]

[13]Similarly, a change in c_R, keeping c_L constant, would imply a direct effect $\left(\frac{\partial}{\partial c_R} E\left(x_R^D\right) < 0\right)$ and an indirect effect $\left(\frac{\partial}{\partial c_R} E\left(x_L^D\right) > 0\right)$.

Figure 1. Relationship between Expected Outputs When the Central Bank Is Moderately Conservative



A Moderately Conservative Central Bank. Delegating monetary policy to a moderately conservative central bank implies that such policy will now be implemented by an authority that is more conservative than party L and less conservative than party R. In other words, there will be an increase in c_L and a decrease in c_R with respect to the initial nondelegation situation ($(c_L, c_R) = (1, 1)$). Notice that, in this case, the direct and indirect effects for expected output under each party work in the same direction. For party L, these effects bring a reduction in expected output and for party R an increase in expected output. Consequently, $E\left(x_L^N\right) > E\left(x_L^D\right)$ and $E\left(x_R^N\right) < E\left(x_R^D\right)$. From lemmas 1 and 2, it follows that $E\left(x_L^N\right) > E\left(x_R^N\right)$ and $E\left(x_L^D\right) > E\left(x_R^D\right)$. Therefore, we obtain $E\left(x_R^N\right) < E\left(x_R^D\right) < E\left(x_L^D\right) < E\left(x_L^N\right)$, as shown in figure 1. In this case the politically induced variance of output is reduced by the presence of an independent and moderately conservative central bank, i.e., $Var_P(x^D) < Var_P(x^N)$.

An Ultraconservative Central Bank. Delegating monetary policy to an ultraconservative central bank implies that such policy will now be implemented by an authority that is more conservative than both parties. Formally, now $c_L > c_R \geq 1$, as there will have been an increase in both c_L and c_R with respect to the initial nondelegation situation ($(c_L, c_R) = (1, 1)$). In this case, the direct and indirect effects on expected outputs work in opposite directions. Notice that the increase in c_L is larger than the increase in c_R. For this reason, the direct effect always dominates for party L and so expected output for this party falls $\left(E\left(x_L^N\right) > E\left(x_L^D\right)\right)$. By contrast, for party R the direct effect might not always dominate. For instance, the indirect effect will be more important for party R when party L is substantially less conservative than party R, or when the central bank is very similar in conservativeness to R. In this last case, i.e., when the indirect effect dominates for party R, expected

Figure 2. Relationship between Outputs When the Central Bank Is Ultraconservative and the Reduction in Expected Output for Party *L* is Larger

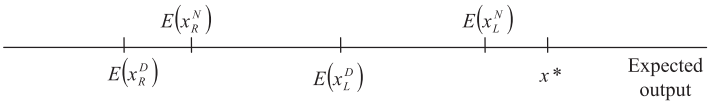


Figure 3. Relationship between Outputs When the Central Bank Is Ultraconservative and the Reduction in Expected Output for Party *L* is Lower

output for this party will increase ($E\left(x_R^N\right) < E\left(x_R^D\right)$). Now, as $E\left(x_L^N\right) > E\left(x_L^D\right)$ and $E\left(x_R^N\right) < E\left(x_R^D\right)$, the analysis related to the comparison of the politically induced variance of output would be identical to the moderately conservative central bank case.

When the direct effect dominates for party R, we can have two possible cases, illustrated in the following two figures. Figure 2 illustrates the case in which the reduction in expected output for party L will be larger and, consequently, $Var_P(x^N) > Var_P(x^D)$.

Figure 3 shows the case in which the reduction in expected output for party L will be smaller than for party R. Therefore, $Var_P(x^N) < Var_P(x^D)$.

The following proposition summarizes these results and identifies the parameter configurations in which the politically induced variance of output is reduced with delegation of monetary policy.

Proposition 3. *(a) By appointing a moderately conservative independent central bank responsible for monetary policy, the politically induced variance of output is reduced. (b) By appointing an ultraconservative independent central bank, the politically induced variance of output is reduced when both parties are concerned enough about public spending stabilization or when the central bank is not too ultraconservative (i.e., when γ is high enough or when δ_{CB} is high enough).*

Proposition 3(b) shows that the politically induced variance of output is reduced with the introduction of an ultraconservative central bank whenever γ is high enough. To understand this result, consider the limiting case when γ converges to infinity. In this case the behavior of both parties would be identical since both would choose the tax rate such that $g_j = g^*$. Consequently, the central bank would select the same inflation rate and, hence, we would expect identical outputs under delegation, resulting in a null politically induced variance of output in this framework. By contrast, under nondelegation we expect different inflation rates due to different preferences between parties, which will generate a strictly positive $Var_P(x^N)$ even though we consider this limiting case ($\gamma \to \infty$).

Proposition 3(b) also indicates that when γ is low enough (but not null) and δ_{CB} is also low, the opposite result is obtained, i.e., $Var_P(x^N) < Var_P(x^D)$.[14] To understand the logic of this result, let us consider the limiting case in which $\delta_{CB} = 0$. Under delegation, the inflation rate would be chosen regardless of output deviations, which is not the case under nondelegation. In addition, for low values of γ, output stabilization is relatively important in the choice of π under nondelegation. This causes outputs to be closer to the target and differ less under nondelegation when γ is low enough and, hence, the politically induced variance of output is increased with delegation of monetary policy.

2.4.2 The Economically Induced Variance of Output

The economically induced variances of output are originated by exogenous shocks. In the model presented here with fiscal policy, these variances when monetary policy is controlled by the government and when it is delegated to an independent central bank are, respectively,

$$Var_E\left(x^N\right) = \left(P\left(\frac{1}{2\delta_L\left(m_L+2\right)}\right)^2 + (1-P)\left(\frac{1}{2\delta_R\left(m_R+2\right)}\right)^2\right)\sigma_\varepsilon^2 \text{ and}$$

[14] If $\gamma = 0$, then $Var_P(x^N) = Var_P(x^D) = 0$.

$$Var_E\left(x^D\right) = \left(P\left(\frac{c_L}{2\delta_L\left(c_L m_L + 2\right)}\right)^2 + (1-P)\left(\frac{c_R}{2\delta_R\left(c_R m_R + 2\right)}\right)^2\right)\sigma_\varepsilon^2.$$

PROPOSITION 4. *The appointment of a moderately conservative independent central bank increases the economically induced variance of output whenever P is large enough. By contrast, the appointment of an ultraconservative central bank always increases the economically induced variance of output.*

This result is in line with the previous literature: appointing an independent central bank more conservative than both parties increases the economically induced variance of output. However, if the central bank's conservativeness is intermediate, then the economically induced variance of output is higher under delegation of monetary policy whenever P is large enough, that is, when the probability of party L—the less inflation-averse party—winning elections is high.

2.5 Output Stabilization

Alesina and Gatti (1995) find that if the volatility of shocks is low enough, then delegation of the conduct of monetary policy reduces the variance of output. This is so because, in this case, the relevant component of the volatility of output is the politically induced variance of output. However, the analysis developed in the previous section allows us to conclude that this result is not robust in our framework. Moreover, we would like to point out that there is another difference between the two models. In Alesina and Gatti (1995), the study of output stabilization coincides with the study of the variance of output. To see this, note that applying the standard statistics theory,

$$E\left(\left(x^i - x^*\right)^2\right) = \left(E\left(x^i - x^*\right)\right)^2 + var(x^i),\ i = N, D.$$

In Alesina and Gatti (1995), $E\left(x^i - x^*\right) = 0$, and hence, $E\left(\left(x^i - x^*\right)^2\right) = var(x^i)$, which indicates that in their model to

study the stabilization of output it suffices to analyze the variance of output. However, when $E\left(x^i - x^*\right) \neq 0$, as in our model, this will not be the case. We can rewrite the output stabilization term as follows:[15]

$$E\left(\left(x^i - x^*\right)^2\right) = P(E(x_L^i - x^*))^2 + (1-P)(E(x_R^i - x^*))^2 + Var_E\left(x^i\right),\ i = N, D.$$

When σ_ε^2 is large enough, the comparison of output stabilization is reduced to the comparison of the economically induced variance of output. By contrast, when σ_ε^2 is low enough, the comparison of output stabilization under delegation and nondelegation involves the analysis of the sum of the first two terms in the previous expression. We know from the analysis carried out previously that the introduction of a moderately conservative independent central bank will lower expected output under a left-wing party and increase it under a right-wing party ($E\left(x_L^N\right) > E\left(x_L^D\right)$ and $E\left(x_R^N\right) < E\left(x_R^D\right)$).[16] Moreover, given that all expected outputs are smaller than the output target, x^*, then $0 < E\left(x^* - x_L^N\right) < E\left(x^* - x_L^D\right)$ and $0 < E\left(x^* - x_R^D\right) < E\left(x^* - x_R^N\right)$. Hence, $\left(E(x_L^D - x^*)\right)^2 > \left(E(x_L^N - x^*)\right)^2$ and $\left(E(x_R^D - x^*)\right)^2 < \left(E(x_R^N - x^*)\right)^2$. Therefore, we can conclude that when σ_ε^2 is low enough and P is high enough, output is more stable under nondelegation, whereas the opposite result is obtained when P is low enough. In other words, whenever the supply shocks are not significant, output stabilization will be more effective without a moderately conservative independent central bank the more likely is the less inflation-averse party to win the elections.

When an ultraconservative central bank is appointed and the indirect effect dominates, we obtain the same result. If the direct effect dominates, then $E\left(x_L^N\right) > E\left(x_L^D\right)$ and $E\left(x_R^N\right) > E\left(x_R^D\right)$. Hence, $0 < E\left(x^* - x_L^N\right) < E\left(x^* - x_L^D\right)$ and $0 < E\left(x^* - x_R^N\right) < E\left(x^* - x_R^D\right)$. Accordingly, $\left(E(x_L^D - x^*)\right)^2 > \left(E(x_L^N - x^*)\right)^2$ and $\left(E(x_R^D - x^*)\right)^2 > \left(E(x_R^N - x^*)\right)^2$. Therefore, we can conclude that,

[15] See lemma A.1 in the appendix.
[16] See figure 1.

in this case, when σ_ε^2 is low enough output is more stable under nondelegation of monetary policy.

3. Robustness

In this section, we test the robustness of the results derived in the benchmark model. We examine two possible variations of our initial model. In subsection 3.1, we extend the model by allowing parties to differ in their target for public spending. In subsection 3.2, we analyze the framework in which authorities choose their policies sequentially, where the fiscal authority acts as the leader and the central bank acts as the follower. As shown below, we conclude that the main results obtained in section 2 are maintained in these other frameworks.

3.1 Different Targets for Public Spending

Next, we will generalize the benchmark model by assuming that both parties differ in their target for public spending. It could be argued that left-wing parties would prefer a higher target for government expenditure than right-wing parties. For this reason, we will assume that party L has a larger government expenditure target than party R: $g_L^* > g_R^*$. In this new setup, we obtain the following inflation rates and outputs under nondelegation:

$$\pi_L^N = \frac{(P + m_R + 1)\, A_L + (1 - P)\, A_R}{\Delta^N} - \frac{\varepsilon}{m_L + 2},$$
$$\pi_R^N = \frac{(1 - P + m_L + 1)\, A_R + P A_L}{\Delta^N} - \frac{\varepsilon}{m_R + 2},$$
$$x_L^N = x^* - \frac{1}{2\delta_L}\pi_L^N, \text{ and } x_R^N = x^* - \frac{1}{2\delta_R}\pi_R^N,$$

where $A_j = g_j^* + w^* + x^*$, $j = L, R$, and under delegation:

$$\pi_L^D = \frac{(c_R m_R + 1 + P)\, A_L + (1 - P)\, A_R}{\Delta^D} - \frac{\varepsilon}{c_L m_L + 2},$$
$$\pi_R^D = \frac{(c_L m_L + 1 + 1 - P)\, A_R + P A_L}{\Delta^D} - \frac{\varepsilon}{c_R m_R + 2},$$
$$x_L^D = x^* - \frac{c_L}{2\delta_L}\pi_L^D, \text{ and } x_R^D = x^* - \frac{c_R}{2\delta_R}\pi_R^D.$$

The following proposition shows that the results obtained in the benchmark model in terms of expected inflation and inflation stabilization still hold in this new setup.

PROPOSITION 5. *(a) By appointing an ultraconservative independent central bank responsible for monetary policy, the expected value of inflation is reduced and a higher degree of inflation stabilization is achieved. (b) By appointing a moderately conservative independent central bank, the expected value of inflation is reduced and a higher degree of inflation stabilization is achieved provided that P is high enough.*[17]

In what follows, we will show that, when parties differ in their views on the size of the government, the results obtained for output in the benchmark model are no longer robust. In particular, how much more concerned about output stabilization party L is with respect to party R will play a crucial role. The following lemma points out that in this general setup expected output might be higher or lower under an L administration.

LEMMA 4. *(a) $E(x_L^N) > E(x_R^N)$ if and only if $\delta_L - \delta_R$ is large enough. (b) $E\left(x_L^D\right) > E\left(x_R^D\right)$ if and only if $\delta_L - \delta_R$ is large enough.*

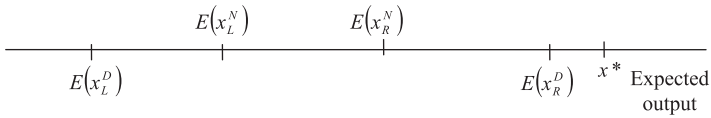
This lemma indicates that the comparison between expected outputs depends on the value of the difference $\delta_L - \delta_R$. Notice that when the difference $\delta_L - \delta_R$ is small, we could expect a lower output when party L is in office. This occurs because, as party L has a higher target for public spending ($g_L^* > g_R^*$), it will be more willing to increase taxes, which will lead to lower output.

Obviously, the discrepancy of results between the two models will also affect the results related to the politically induced variance of output, as shown in the following proposition.

PROPOSITION 6. *(a) By appointing a moderately conservative independent central bank responsible for monetary policy, the politically*

[17]The results derived in propositions 1 and 5 also hold in an extension of the benchmark model in which the assumption that all policymakers have identical output targets is relaxed by considering that the left-wing party has a higher target.

Figure 4. Expected Outputs When $g_L^* > g_R^*, \delta_L - \delta_R$ Is Low and the Central Bank Is Moderately Conservative



induced variance of output is reduced provided that the difference $\delta_L - \delta_R$ is large enough. (b) By appointing an ultraconservative independent central bank, the politically induced variance of output is reduced when $\delta_L - \delta_R$ is large enough for high values of $g_L^* - g_R^*$. For low values of $g_L^* - g_R^*$, it is required that $\delta_L - \delta_R$ is large enough and that (i) both parties are concerned enough about public spending stabilization or (ii) the central bank is not too ultraconservative.

Unlike the benchmark model, the introduction of a moderately conservative independent central bank increases the politically induced variance of output whenever $\delta_L - \delta_R$ is low enough. Remember that delegating monetary policy to a moderately conservative central bank decreases expected output for party L and increases expected output for party R. As expected output for party L will be smaller than party R's under nondelegation, the overall effect is that $E\left(x_L^D\right) < E\left(x_L^N\right) < E\left(x_R^N\right) < E\left(x_R^D\right)$. Hence, the politically induced variance of output with a moderately conservative independent central bank is higher than with no independent monetary policy, i.e., $Var_P(x^N) < Var_P(x^D)$. We show this situation in figure 4.

Furthermore, when party L places much more weight on output stabilization ($\delta_L - \delta_R$ is large enough), lemma 4 states that $E\left(x_L^N\right) > E\left(x_R^N\right)$ and $E\left(x_L^D\right) > E\left(x_R^D\right)$, as in the benchmark model, which results in the same comparison of the politically induced variance of output, i.e., $Var_P(x^D) < Var_P(x^N)$.

Proposition 6(b) shows that the introduction of an ultraconservative independent central bank will increase the politically induced variance of output whenever $\delta_L - \delta_R$ is low enough. In this case, expected output under party L will always be lower than under party R ($E\left(x_R^N\right) > E\left(x_L^N\right)$ and $E\left(x_R^D\right) > E\left(x_L^D\right)$). The fact

Figure 5. Expected Outputs When $g_L^* > g_R^*, \delta_L - \delta_R$ Is Low and the Central Bank Is Ultraconservative

that party L has a higher target for public spending will bring in a stronger response from the ultraconservative central bank to reduce inflation and, thus, it will result in a larger reduction in expected output for party L,[18] as shown in figure 5. Consequently, $Var_P(x^N) < Var_P(x^D)$.

By contrast, when $\delta_L - \delta_R$ is large enough, the comparison between the politically induced variances of output depends on the difference $g_L^* - g_R^*$. For low values of this difference, we obtain the same results as in the benchmark case, but for high values of $g_L^* - g_R^*$ we find that $E\left(x_L^N\right) - E\left(x_R^N\right) > E\left(x_L^D\right) - E\left(x_R^D\right)$. As both sides of the previous inequality are positive whenever $\delta_L - \delta_R$ is large enough, it holds that $Var_P(x^N) > Var_P(x^D)$.

In relation to the economically induced variance of output, notice that the terms that multiply the supply shock in the expressions for output under delegation and nondelegation derived in this extension and in the benchmark model coincide. Lemma A.1 implies that we obtain the same results in this setup as in the basic framework. Therefore, proposition 4 applies in this new setup.

Additionally, it can be shown that $E\left(x_L^N\right) > E\left(x_L^D\right)$ is always satisfied, while $E\left(x_R^N\right) < E\left(x_R^D\right)$ if the central bank is moderately conservative. As a result, the analysis included in subsection 2.5 also applies in this extension and, therefore, we can conclude that the results related to output derived in the benchmark model are robust.

[18] If δ_L and δ_R are similar, the parties' behavior basically differs because of the different target for public spending. As the central bank chooses the inflation rate without taking into account public spending, under nondelegation we expect a large change in output for the party that chooses inflation with the largest target for public spending (party L), i.e., $E\left(x_L^N\right) - E\left(x_L^D\right) > E\left(x_R^N\right) - E\left(x_R^D\right)$.

3.2 Fiscal Leadership

The assumption that the government and the central bank choose their policies simultaneously is rather unrealistic. Taxes tend to be changed when the yearly budget is approved or when there is a change in the government. Money supply or interest rates, on the other hand, can be changed throughout the year. Thus, changing taxes is in general a more time-consuming process than adjusting the stance of monetary policy. Therefore, a more appropriate description of fiscal–monetary interactions would involve a leader–follower game, in which the fiscal authority acts as the leader and the monetary authority acts as the follower. Therefore, in this case the fiscal authority would choose the tax rate anticipating the response of the central bank to its action.

The following lemma shows that expected output will also be higher under an L administration when the government acts as a Stackelberg leader.[19]

Lemma 5. $E\left(x_L^{D,S}\right) > E\left(x_R^{D,S}\right)$.

Lemma 5 also implies that the politically induced variance of output is not automatically eliminated by delegating the monetary policy to an independent central bank in this sequential game.

The next two propositions show that the effects on inflation and output of appointing an independent central bank are not altered when introducing sequentiality in policy actions.

Proposition 7. *(a) By appointing an ultraconservative independent central bank responsible for monetary policy, the expected value of inflation is reduced and a higher degree of inflation stabilization is achieved. (b) By appointing a moderately conservative independent central bank, the expected value of inflation is reduced and a higher degree of inflation stabilization is achieved provided that P is high enough.*

[19]The superscript S indicates that we are considering a Stackelberg game. A detailed derivation of the optimal inflation rates and outputs under this new setup is given in proposition C.1.

Proposition 8. *(a) The appointment of a moderately conservative independent central bank reduces the politically induced variance of output. By contrast, the appointment of an ultraconservative central bank reduces the politically induced variance of output when both parties are concerned enough about public spending stabilization or when the central bank is not too ultraconservative (i.e., when γ is high enough or when δ_{CB} is high enough). (b) The appointment of a moderately conservative independent central bank increases the economically induced variance of output whenever P is large enough. By contrast, the appointment of an ultraconservative central bank always increases the economically induced variance of output.*

Finally, in this new setup it can also be shown that $E\left(x_L^N\right) > E\left(x_L^{D,S}\right)$ is always satisfied, while $E\left(x_R^N\right) < E\left(x_R^{D,S}\right)$ if the central bank is moderately conservative. Applying the same reasoning as in the previous extension of the benchmark case, we conclude that the results related to output stabilization derived in the benchmark model are robust.

4. Conclusions

The analysis presented in this article has shown that the extent to which an independent central bank is successful in attaining price and output stability depends on the degree of conservativeness of the central bank in relation to the political parties, the private sector's expectations on which party will win, and the level of economic uncertainty.

Alesina and Stella (2010) point out that in an economic crisis the level of economic uncertainty is high. In such a case, according to the model presented here, we have shown that the appointment of an ultraconservative central bank always increases output instability. By contrast, the appointment of a moderately conservative independent central bank will increase output instability if the probability of the less inflation-averse party winning the elections is high enough.

Further, when there is little economic uncertainty, Alesina and Gatti's (1995) result that delegation of the conduct of monetary policy increases output stabilization is not always valid. In particular, we have shown that a moderately conservative central bank

reduces output stabilization if the probability of the less inflation-averse party winning the elections is high enough. The same result applies for an ultraconservative central bank when the left-wing party is substantially less conservative than the right-wing party, or when the central bank is very similar in conservativeness to the right-wing party. In other cases, an ultraconservative central bank always reduces output stability.

Focusing on the European case, the Maastricht Treaty determined that the European Central Bank should primarily be concerned with price stability in the euro area. Our analysis suggests that the creation of such an ultraconservative central bank has resulted in a lower and more stable inflation, but likely at a cost of more output instability, even in the case when there is little economic uncertainty. In particular, Martínez-Martín, Saiz, and Stoevsky (2018) show how growth volatility across the euro-area countries was substantially higher than for the G-7 countries in the period after the global financial crisis (2009 to 2012).[20]

We have checked the robustness of our results considering some variations of our initial model. There are some other possible interesting avenues of research. For instance, a natural extension would be to endogenize the probability of a party being elected. In this potentially dynamic setting, public debt could also be introduced in order to consider how the incumbent party could affect the probability of being elected.

Finally, we would like to point out that our key results depend on the manner in which taxes are introduced in Alesina and Gatti's model. The model presented in this article follows the related literature where taxes lower aggregate supply. However, taxes could alternatively be incorporated through aggregate demand. In this case, the inflation rate and the level of output chosen by a central bank would not depend on the party in office, which would be aligned with Alesina and Gatti's (1995) result.[21]

[20]Further, the authors mention that Ireland was excluded "to avoid distortions in the analysis caused by the high volatility of Irish GDP."

[21]We would like to thank an anonymous referee for pointing out this issue.

Appendix A

We initially derive the expressions for output and public spending when government j is in office. To ease the notation, we omit the subscript j in the following two proofs.

Derivation of Expression (1)

Output of a representative firm is given by $X = L^{\lambda}\exp(\varepsilon/2)$, where X denotes the real output, L represents labor, λ indicates the output elasticity, and ε represents a supply shock. We assume that ε has zero mean and variance σ_{ε}^{2}. Distortionary taxes are levied on production. The firm maximizes profit, given by $(1-\tau)PL^{\lambda}\exp(\varepsilon/2) - WL$, where τ denotes the tax rate on total revenue of firms, P represents the price level, and W is the nominal wage. Solving for the firm's labor demand, assuming it can hire the labor it demands at the given nominal wage, taking logs, we have $x = a\,(p-\tau-w) + b + \frac{\varepsilon}{2(1-\lambda)}$, where lowercase letters denote logs of nominal variables, $a = \lambda/(1-\lambda)$, $b = \lambda\ln\lambda/(1-\lambda)$, and $\ln(1-\tau) \approx -\tau$. Following Debelle and Fischer (1994), for simplicity, we set $\lambda = 0.5$, so that $a = 1$ and we approximate $\ln\lambda$ to 0. Hence, $x = p - \tau - w + \varepsilon$.

In addition, following Alesina and Tabellini (1987), Debelle and Fischer (1994), and Beetsma and Bovenberg (2001), among others, we assume that workers are represented by a centralized trade union which seeks to minimize deviations of the real wage rate from a particular target w^{*}, hence it sets the nominal wage (in logs) to achieve the target w^{*}. The trade union chooses the nominal wage in advance of the actions of the two policymakers, but knowing their objective functions, i.e., it minimizes the objective function: $E\left((w-p-w^{*})^{2}\right)\Big/2$. The first-order condition of this optimization problem immediately yields $w = p^{e} + w^{*}$. Therefore, $x = p - p^{e} - \tau - w^{*} + \varepsilon$. Finally, approximating $p - p^{e}$ by $\pi - \pi^{e}$, where π represents inflation rate and π^{e} expected inflation rate, we get the aggregate supply equation of the model, i.e., $x = \pi - \pi^{e} - \tau - w^{*} + \varepsilon$. ■

Derivation of Expression (2)

The government's budget constraint in nominal terms at t is given by $P_tG_t = \tau P_tX_t + M_t - M_{t-1}$, where G_t denotes the public spending and M_t the nominal money supply at t. Following Canzoneri (1985), money demand depends only on an output level that is independent of fiscal policy (taxes), $M_t = P_t\bar{X}$. Dividing the government's budget constraint by nominal income, $P_t\bar{X}$, yields $g_t = \tau\frac{X_t}{\bar{X}} + \frac{M_t - M_{t-1}}{P_t\bar{X}}$, where g_t is public spending as a share of the (nondistortionary) output $\bar{X}$. Taking into account the money demand function, whenever X_t is close to $\bar{X}$ and approximating $\frac{P_t - P_{t-1}}{P_t}$ to π_t (as in Alesina and Tabellini 1987 and Dimakou 2013), we get the government budget constraint at t, i.e., $g_t = \tau + \pi_t$. ■

In the following two propositions, we derive the policies chosen by the two parties, if in office, under nondelegation and under delegation in the benchmark model, where parties differ only in the relative weight assigned to output stabilization.

Proposition A.1. *The policies chosen by the two parties, if in office, under nondelegation are given by*

$$\pi_L^N = \frac{m_R + 2}{\Delta^N}A - \frac{\varepsilon}{m_L + 2},\ \pi_R^N = \frac{m_L + 2}{\Delta^N}A - \frac{\varepsilon}{m_R + 2},$$
$$\tau_L^N = g^* - \left(1 + \frac{1}{2\gamma}\right)\pi_L^N,\ \textit{and}\ \tau_R^N = g^* - \left(1 + \frac{1}{2\gamma}\right)\pi_R^N,$$

where $\Delta^N = P\left(m_R + 2\right)\left(m_L + 1\right) + \left(1 - P\right)\left(m_R + 1\right)\left(m_L + 2\right)$ *and* $A = g^* + w^* + x^*$.

Proof of Proposition A.1. Under nondelegation,[22] the party in office, denoted by j, chooses π_j and τ_j in order to solve the following problem:

$$\underset{\pi_j,\tau_j}{min} V_{Gj} = \frac{1}{2}\left(\pi_j^2 + \delta_j\left(x_j - x^*\right)^2 + \gamma(g_j - g^*)^2\right).$$

[22]To ease the notation, we drop the superscript N and D in the proofs of propositions A.1 and A.2, respectively.

The first-order conditions (FOC) of this problem are given by[23]

$$\frac{\partial}{\partial \pi_j} V_{Gj} = \pi_j + \delta_j \left(x_j - x^*\right) + \gamma(g_j - g^*) = 0 \text{ and}$$
$$\frac{\partial}{\partial \tau_j} V_{Gj} = -\delta_j \left(x_j - x^*\right) + \gamma(g_j - g^*) = 0.$$

Using expressions (1) and (2) in the previous two equalities, we get

$$\pi_j = \frac{\pi^e + A - \varepsilon}{m_j + 2} \text{ and} \tag{A.1}$$

$$\tau_j = g^* - \frac{\delta_j \left(2\gamma + 1\right)}{\gamma + \delta_j + 4\gamma\delta_j} \left(\pi^e + A - \varepsilon\right), \tag{A.2}$$

where $m_j = \frac{\frac{1}{\delta_j} + \frac{1}{\gamma}}{2}$ and $A = g^* + w^* + x^*$. Rewriting (A.1) for the two parties, we have $\pi_L = \frac{\pi^e + A - \varepsilon}{m_L + 2}$ and $\pi_R = \frac{\pi^e + A - \varepsilon}{m_R + 2}$. Moreover, recall that $\pi^e = PE\left(\pi_L\right) + (1 - P)E\left(\pi_R\right)$. Taking expectations in the previous expressions and solving for π^e, we get

$$\pi^e = \frac{\frac{P}{m_L + 2} + \frac{1-P}{m_R + 2}}{1 - \left(\frac{P}{m_L + 2} + \frac{1-P}{m_R + 2}\right)} A. \tag{A.3}$$

Substituting this expression into (A.1) and (A.2) for $j = L, R$, and after some algebra, we obtain the expressions for π_L, π_R, τ_L, and τ_R. ■

Proposition A.2. *Under delegation, the policies chosen by the central bank and the party, if in office, are given by*

$$\pi_L^D = \frac{c_R m_R + 2}{\Delta^D} A - \frac{\varepsilon}{c_L m_L + 2}, \; \pi_R^D = \frac{c_L m_L + 2}{\Delta^D} A - \frac{\varepsilon}{c_R m_R + 2},$$
$$\tau_L^D = g^* - \left(1 + \frac{c_L}{2\gamma}\right) \pi_L^D, \text{ and } \tau_R^D = g^* - \left(1 + \frac{c_R}{2\gamma}\right) \pi_R^D,$$

[23]Direct computations yield that the objective function is strictly convex. Therefore, the first-order conditions are necessary and sufficient to obtain a minimum. The same comment applies to the remaining optimization problems.

where

$$\Delta^D = P\left(c_L m_L + 1\right)\left(c_R m_R + 2\right) + \left(1 - P\right)\left(c_R m_R + 1\right)\left(c_L m_L + 2\right).$$

Proof of Proposition A.2. Suppose that party j is in office. Under delegation, the central bank chooses π_j in order to solve the following problem:

$$\underset{\pi_j}{min} V_{CB} = \frac{1}{2}\left(\pi_j^2 + \delta_{CB}\left(x_j - x^*\right)^2\right).$$

The FOC of this problem is given by $\frac{\partial}{\partial \pi_j} V_{CB} = \pi_j + \delta_{CB}(x_j - x^*) = 0$.

In this setup the fiscal authority chooses τ_j in order to solve the following problem:

$$\underset{\tau_j}{min} V_{G_j} = \frac{1}{2}\left(\pi_j^2 + \delta_j\left(x_j - x^*\right)^2 + \gamma(g_j - g^*)^2\right).$$

The FOC of this problem is given by $\frac{\partial}{\partial \tau_j} V_{G_j} = -\delta_j\left(x_j - x^*\right) + \gamma(g_j - g^*) = 0$. Using expressions (1) and (2) in the FOC of the authorities' problems, it follows that

$$\pi_j = \frac{\delta_{CB}\gamma}{\gamma + \delta_j + 2\gamma\delta_{CB}}\left(\pi^e + A - \varepsilon\right) \text{ and} \tag{A.4}$$

$$\tau_j = g^* - \frac{\delta_j + \gamma\delta_{CB}}{\gamma + \delta_j + 2\gamma\delta_{CB}}\left(\pi^e + A - \varepsilon\right). \tag{A.5}$$

Using them in the expression for π^e and solving for π^e, we get

$$\pi^e = \frac{\delta_{CB}\gamma\left(\frac{P}{\gamma+\delta_L+2\gamma\delta_{CB}} + \frac{1-P}{\gamma+\delta_R+2\gamma\delta_{CB}}\right)A}{1 - \delta_{CB}\gamma\left(\frac{P}{\gamma+\delta_L+2\gamma\delta_{CB}} + \frac{1-P}{\gamma+\delta_R+2\gamma\delta_{CB}}\right)}.$$

Substituting this expression into (A.4) and (A.5) for $j = L, R$, and after some algebra, we obtain the desired expressions. ■

Proof of Lemma 1

Combining the FOC of the optimization problem given in the proof of proposition A.1, (6) and (7) are derived. Using the expressions of π_L^N and π_R^N given in the statement of proposition A.1, we get

$$x_L^N = x^* - \frac{1}{2\delta_L}\left(\frac{m_R+2}{\Delta^N}A - \frac{\varepsilon}{m_L+2}\right) \text{ and} \tag{A.6}$$

$$x_R^N = x^* - \frac{1}{2\delta_R}\left(\frac{m_L+2}{\Delta^N}A - \frac{\varepsilon}{m_R+2}\right). \tag{A.7}$$

Taking expectations, we have $E\left(x_L^N\right) = x^* - \frac{m_R+2}{2\delta_L\Delta^N}A$ and $E\left(x_R^N\right) = x^* - \frac{m_L+2}{2\delta_R\Delta^N}A$. Using the expressions of m_L and m_R, we have $E\left(x_L^N\right) - E\left(x_R^N\right) = \frac{(4\gamma+1)(\delta_L-\delta_R)}{4\gamma\delta_L\delta_R\Delta^N}A$. As $\delta_L > \delta_R$, we conclude that $E\left(x_L^N\right) > E\left(x_R^N\right)$. ■

Proof of Lemma 2

From the FOC of the optimization problem of the central bank given in the proof of proposition A.2, we have (11) and (12). Using the expressions of π_L^D and π_R^D given in the statement of proposition A.2, it follows that

$$x_L^D = x^* - \frac{c_L}{2\delta_L}\left(\frac{c_R m_R+2}{\Delta^D}A - \frac{\varepsilon}{c_L m_L+2}\right) \text{ and} \tag{A.8}$$

$$x_R^D = x^* - \frac{c_R}{2\delta_R}\left(\frac{c_L m_L+2}{\Delta^D}A - \frac{\varepsilon}{c_R m_R+2}\right). \tag{A.9}$$

Taking expectations, we have

$$E\left(x_L^D\right) = x^* - \frac{c_L\left(c_R m_R+2\right)}{2\delta_L\Delta^D}A \text{ and} \tag{A.10}$$

$$E\left(x_R^D\right) = x^* - \frac{c_R\left(c_L m_L+2\right)}{2\delta_R\Delta^D}A. \tag{A.11}$$

From the expressions of m_L, m_R, c_L, and c_R, $E\left(x_L^D\right) - E\left(x_R^D\right) = \frac{\delta_L-\delta_R}{\delta_{CB}^2\gamma\Delta^D}A$, which implies $E\left(x_L^D\right) > E\left(x_R^D\right)$ since $\delta_L > \delta_R$. ■

Next, we derive a lemma which will be useful to prove some of the results that follow.

Lemma A.1. *Consider a random variable z that, conditional on the realization of the shock, takes two possible values given by $z_L = E(z_L) + F_L\varepsilon$ and $z_R = E(z_R) + F_R\varepsilon$. Then, the politically induced*

variance of z is given by $Var_P(z) = P(1-P)(E(z_L)-E(z_R))^2$, and the economically induced variance of z by $Var_E(z) = (P(F_L)^2 + (1-P)(F_R)^2)\sigma_\varepsilon^2$. Moreover,

$$E\left(z^2\right) = P(E(z_L))^2 + (1-P)(E(z_R))^2 + \left(P\left(F_L\right)^2 + (1-P)\left(F_R\right)^2\right)\sigma_\varepsilon^2. \tag{A.12}$$

Proof of Lemma A.1. Notice that $Var(z) = PE((z_L - E(z))^2) + (1-P)E((z_R - E(z))^2)$. As $E((z_L - E(z))^2) = (1-P)^2(E(z_L) - E(z_R))^2 + (F_L)^2\sigma_\varepsilon^2$ and $E((z_R - E(z))^2) = P^2(E(z_L) - E(z_R))^2 + (F_R)^2\sigma_\varepsilon^2$, we have $Var(z) = P(1-P)(E(z_L) - E(z_R))^2 + (P(F_L)^2 + (1-P)(F_R)^2)\sigma_\varepsilon^2$. The first term corresponds to the politically induced variance of z, whereas the second term corresponds to the economically induced variance of z. Finally, expression (A.12) follows from the fact that $E\left(z^2\right) = \left(E\left(z\right)\right)^2 + Var\left(z\right)$. ■

Proof of Proposition 1

Direct computations yield

$$E\left(\pi^N\right) = \frac{P\left(m_R+2\right) + (1-P)\left(m_L+2\right)}{\Delta^N}A \text{ and}$$

$$E\left(\pi^D\right) = \frac{P\left(c_R m_R+2\right) + (1-P)\left(c_L m_L+2\right)}{\Delta^D}A.$$

In addition, applying lemma A.1 for $z = \pi^N$ and $z = \pi^D$, it follows that

$$E\left((\pi^N)^2\right) = \frac{P\left(m_R+2\right)^2 + (1-P)\left(m_L+2\right)^2}{\left(\Delta^N\right)^2}A^2 + \left(P\left(\frac{1}{m_L+2}\right)^2 + (1-P)\left(\frac{1}{m_R+2}\right)^2\right)\sigma_\varepsilon^2 \text{ and}$$

$$E\left((\pi^D)^2\right) = \frac{P\left(c_R m_R+2\right)^2 + (1-P)\left(c_L m_L+2\right)^2}{\left(\Delta^D\right)^2}A^2 + \left(P\left(\frac{1}{c_L m_L+2}\right)^2 + (1-P)\left(\frac{1}{c_R m_R+2}\right)^2\right)\sigma_\varepsilon^2.$$

(a) Let $f(c_L, c_R) = E\left(\pi^D\right)$ and $g(c_L, c_R) = E\left((\pi^D)^2\right)$. Notice that $f(c_L, c_R)$ and $g(c_L, c_R)$ are decreasing functions in c_L and c_R. Moreover, $f(1,1) = E\left(\pi^N\right)$ and $g(1,1) = E\left((\pi^N)^2\right)$. The combination of these results allows us to conclude that $E\left(\pi^N\right) > E\left(\pi^D\right)$ and $E\left((\pi^N)^2\right) > E\left((\pi^D)^2\right)$ whenever $c_L > c_R \geq 1$.

(b) Suppose now that $c_L > 1 > c_R$. First, we focus on the comparison of the expected inflation. Let $h(P) = E\left(\pi^N\right) - E\left(\pi^D\right)$. Differentiating,

$$\frac{\partial}{\partial P}h(P) = \frac{(m_L+2)(m_R+2)(m_R-m_L)}{(\Delta^N)^2}A - \frac{(c_L m_L+2)(c_R m_R+2)(c_R m_R - c_L m_L)}{(\Delta^D)^2}A.$$

Direct computations yield $\frac{\partial}{\partial c_L}(\frac{\partial}{\partial P}h(P)) > 0$ and $\frac{\partial}{\partial c_R}(\frac{\partial}{\partial P}h(P)) < 0$. Hence, $\frac{\partial}{\partial P}h(P) > \frac{\partial}{\partial P}h(P)|_{\substack{c_L=1\\c_R=1}} = 0$ since $c_L > 1 > c_R$. Therefore, $h(P)$ is an increasing function in P. Moreover, $h(1) > 0$ and $h(0) < 0$ whenever $c_L > 1 > c_R$. This implies that there exists a unique value $\overline{P}$ such that $h(P) > 0$ (or equivalently, $E(\pi^N) > E(\pi^D)$) if and only if $P > \overline{P}$.

In relation to the comparison of the term related to inflation stabilization, notice that as $c_L > 1 > c_R$, $E((\pi^N)^2)|_{P=1} > E((\pi^D)^2)|_{P=1}$ and $E((\pi^N)^2)|_{P=0} < E((\pi^D)^2)|_{P=0}$. Moreover, direct computations yield: (1) $E((\pi^N)^2)$ increases in P since $m_R > m_L$, and (2) $E((\pi^D)^2)$ decreases in P since $c_R m_R < c_L m_L$. Therefore, we can conclude that there exists a value $\overline{P}$ such that $E((\pi^D)^2) < E((\pi^N)^2)$ if and only if $P > \overline{P}$. ■

Proof of Lemma 3

Differentiating (A.10) and (A.11), we have the results stated in the statement of this lemma. ■

Proof of Proposition 3

Direct computations yield

$$E\left(x_L^N\right) - E\left(x_R^N\right) = \frac{(4\gamma+1)(\delta_L-\delta_R)\gamma}{\begin{pmatrix} P(\gamma+\delta_L+2\gamma\delta_L)(\gamma+\delta_R+4\gamma\delta_R) \\ +(1-P)(\gamma+\delta_R+2\gamma\delta_R)(\gamma+\delta_L+4\gamma\delta_L) \end{pmatrix}} A$$

and

$$E(x_L^D) - E(x_R^D) = \frac{(\delta_L-\delta_R)\gamma}{\begin{pmatrix} P(\gamma+\delta_L+\gamma\delta_{CB})(\gamma+\delta_R+2\gamma\delta_{CB}) \\ +(1-P)(\gamma+\delta_R+\gamma\delta_{CB})(\gamma+\delta_L+2\gamma\delta_{CB}) \end{pmatrix}} A.$$

Using lemma 1 and lemma 2, we have that $Var_P(x^N) > Var_P(x^D)$ if and only if $E\left(x_L^N\right) - E\left(x_R^N\right) > E(x_L^D) - E(x_R^D)$, which is equivalent to $g(\delta_{CB}) > 0$, with

$$g(\delta_{CB}) = 2\gamma^2\delta_{CB}^2 + \delta_{CB}\left(\gamma(\gamma+\delta_L) + 2\gamma(\gamma+\delta_R) + P\gamma(\delta_L-\delta_R)\right) + 2\gamma\frac{2\gamma^2 + (\delta_R + P\delta_L - P\delta_R - 4\delta_L\delta_R)\gamma - \delta_L\delta_R}{4\gamma+1}.$$

Note that g is increasing in δ_{CB}. Next, we distinguish two cases:

Case 1: The central bank is moderately conservative ($2\delta_R < \delta_{CB} < 2\delta_L$). Combining the monotonicity property of $g(\delta_{CB})$ and the fact that $g(2\delta_R) > 0$, we have that $g(\delta_{CB}) > 0$ whenever $2\delta_R < \delta_{CB} < 2\delta_L$ and, hence, $Var_P(x^N) > Var_P(x^D)$.

Case 2: The central bank is ultraconservative ($\delta_{CB} \leq 2\delta_R$). Then, we consider two subcases:

- *Subcase 2.1:* $g(0) \geq 0$. In this case $2\gamma^2 + \gamma(\delta_R + P\delta_L - P\delta_R - 4\delta_L\delta_R) - \delta_L\delta_R \geq 0$. Descartes' rule tells us that there exists a unique value of γ, denoted by $\overline{\gamma}$, such that the previous inequality is satisfied whenever $\gamma \geq \overline{\gamma}$. Combining the monotonicity property of $g(\delta_{CB})$ and the fact that $g(0) \geq 0$, we have that $g(\delta_{CB}) > 0$ whenever $\delta_{CB} > 0$ and,

hence, $Var_P(x^N) > Var_P(x^D)$. Therefore, we conclude that if $\gamma \geq \overline{\gamma}$, then $Var_P(x^N) > Var_P(x^D)$.

- *Subcase 2.2:* $g(0) < 0$ (or, equivalently, $\gamma < \overline{\gamma}$). In this case, the monotonicity property of $g(\delta_{CB})$ implies that there exists a unique value of δ_{CB}, denoted by $\overline{\delta_{CB}}$, such that $g(\delta_{CB}) > 0$ whenever $\delta_{CB} > \overline{\delta_{CB}}$. Consequently, $Var_P(x^N) > Var_P(x^D)$ if and only if $\delta_{CB} > \overline{\delta_{CB}}$. ■

Proof of Proposition 4

Combining (A.6), (A.7), and lemma A.1,

$$Var_E\left(x^N\right) = \left(P\left(\frac{1}{2\delta_L\left(m_L+2\right)}\right)^2 + (1-P)\left(\frac{1}{2\delta_R\left(m_R+2\right)}\right)^2\right)\sigma_\varepsilon^2.$$

Analogously, combining (A.8), (A.9), and lemma A.1,

$$Var_E\left(x^D\right) = \left(P\left(\frac{c_L}{2\delta_L\left(c_L m_L+2\right)}\right)^2 + (1-P)\left(\frac{c_R}{2\delta_R\left(c_R m_R+2\right)}\right)^2\right)\sigma_\varepsilon^2.$$

Hence, $Var_E\left(x^N\right) - Var_E\left(x^D\right)$ is a linear function in P. Next we distinguish two cases:

Case 1: $c_R < 1 < c_L$. It is easy to see that $Var_E(x^N)|_{P=1} - Var_E(x^D)|_{P=1} < 0$ and that $Var_E(x^N)|_{P=0} - Var_E(x^D)|_{P=0} > 0$ whenever $c_R < 1 < c_L$. Hence, we can conclude that there exists a value of P, denoted by $\overline{\overline{P}}$, such that $Var_E(x^D) > Var_E(x^N)$ if and only if $P > \overline{\overline{P}}$.

Case 2: $c_L > c_R \geq 1$. Direct computations yield that $Var_E(x^D) > Var_E(x^N)$ whenever $c_L > c_R \geq 1$. ■

Appendix B

In the following two propositions, we derive the policies chosen by the two parties, if in office, under nondelegation and under delegation in the generalized model, where parties differ both in the

relative weight assigned to output stabilization and in their target for public spending.

Proposition B.1. *The policies chosen by the two parties, if in office, under nondelegation are given by*

$$\pi_L^N = \frac{(P + m_R + 1)\, A_L + (1 - P)\, A_R}{\Delta^N} - \frac{\varepsilon}{m_L + 2},$$
$$\pi_R^N = \frac{(1 - P + m_L + 1)\, A_R + P A_L}{\Delta^N} - \frac{\varepsilon}{m_R + 2},$$
$$\tau_L^N = g_L^* - \left(1 + \frac{1}{2\gamma}\right) \pi_L^N, \text{ and } \tau_R^N = g_R^* - \left(1 + \frac{1}{2\gamma}\right) \pi_R^N,$$

where $\Delta^N = P\,(m_R + 2)\,(m_L + 1) + (1 - P)\,(m_R + 1)\,(m_L + 2)$ *and* $A_j = g_j^* + w^* + x^*$, $j = L, R$.

Proof of Proposition B.1. Under nondelegation, the party in office, denoted by j, chooses π_j and τ_j in order to solve the following problem:

$$\underset{\pi_j, \tau_j}{min} V_{Gj} = \frac{1}{2}\left(\pi_j^2 + \delta_j\,(x_j - x^*)^2 + \gamma(g_j - g_j^*)^2\right).$$

The FOCs of this problem are given by

$$\frac{\partial}{\partial \pi_j} V_{Gj} = \pi_j + \delta_j\,(x_j - x^*) + \gamma(g_j - g_j^*) = 0 \text{ and}$$
$$\frac{\partial}{\partial \tau_j} V_{Gj} = -\delta_j\,(x_j - x^*) + \gamma(g_j - g_j^*) = 0.$$

Using expressions (1) and (2) in the previous two equalities, we get

$$\pi_j = \frac{1}{m_j + 2}\,(\pi^e + A_j - \varepsilon) \text{ and} \tag{B.1}$$

$$\tau_j = g_j^* - \frac{\delta_j\,(2\gamma + 1)}{\gamma + \delta_j + 4\gamma\delta_j}\,(\pi^e + A_j - \varepsilon), \tag{B.2}$$

where $A_j = g_j^* + w^* + x^*$. Rewriting (B.1) for the two parties, we have $\pi_L = \frac{\pi^e + A_L - \varepsilon}{m_L + 2}$ and $\pi_R = \frac{\pi^e + A_R - \varepsilon}{m_R + 2}$. Moreover, recall that

$\pi^e = PE(\pi_L) + (1-P)E(\pi_R)$. Taking expectations in the previous expressions and solving for π^e, we get

$$\pi^e = \frac{\frac{P}{m_L+2}A_L + \frac{1-P}{m_R+2}A_R}{1 - \left(P\frac{1}{m_L+2} + (1-P)\frac{1}{m_R+2}\right)}.$$

Substituting this expression into (B.1) and (B.2) for $j = L, R$, and after some algebra, we obtain the expressions for π_L, π_R, τ_L, and τ_R. ■

Proposition B.2. *Under delegation, the policies chosen by the central bank and the party, if in office, are given by*

$$\pi_L^D = \frac{(c_R m_R + 1 + P)\, A_L + (1-P)\, A_R}{\Delta^D} - \frac{\varepsilon}{c_L m_L + 2},$$

$$\pi_R^D = \frac{(c_L m_L + 1 + 1 - P)\, A_R + P A_L}{\Delta^D} - \frac{\varepsilon}{c_R m_R + 2},$$

$$\tau_L^D = g_L^* - \left(1 + \frac{c_L}{2\gamma}\right)\pi_L^D, \text{ and } \tau_R^D = g_R^* - \left(1 + \frac{c_R}{2\gamma}\right)\pi_R^D,$$

where $\Delta^D = P(c_L m_L + 1)(c_R m_R + 2) + (1-P)(c_R m_R + 1)(c_L m_L + 2)$.

Proof of Proposition B.2. Suppose that party j is in office. Under delegation, the central bank chooses π_j in order to solve the following problem:

$$\min_{\pi_j} V_{CB} = \frac{1}{2}\left(\pi_j^2 + \delta_{CB}\left(x_j - x^*\right)^2\right).$$

The FOC of this problem is given by $\frac{\partial}{\partial \pi_j} V_{CB} = \pi_j + \delta_{CB}(x_j - x^*) = 0$.

In this setup the fiscal authority chooses τ_j in order to solve the following problem:

$$\min_{\tau_j} V_{G_j} = \frac{1}{2}\left(\pi_j^2 + \delta_j\left(x_j - x^*\right)^2 + \gamma(g_j - g_j^*)^2\right).$$

The FOC of this problem is given by $\frac{\partial}{\partial \tau_j} V_{G_j} = -\delta_j\left(x_j - x^*\right) + \gamma(g_j - g_j^*) = 0$. Using expressions (1) and (2) in the FOC of the authorities' problems, it follows that

$$\pi_j = \frac{\gamma\delta_{CB}}{\gamma + \delta_j + 2\gamma\delta_{CB}}\left(\pi^e + A_j - \varepsilon\right) \text{ and} \tag{B.3}$$

$$\tau_j = g_j^* - \frac{\delta_j + \gamma\delta_{CB}}{\gamma + \delta_j + 2\gamma\delta_{CB}} \left(\pi^e + A_j - \varepsilon\right). \tag{B.4}$$

Using them in the expression for π^e and solving for π^e, we get

$$\pi^e = \frac{\delta_{CB}\gamma \left(\frac{P}{\gamma+\delta_L+2\gamma\delta_{CB}} A_L + \frac{1-P}{\gamma+\delta_R+2\gamma\delta_{CB}} A_R\right)}{1 - \delta_{CB}\gamma \left(\frac{P}{\gamma+\delta_L+2\gamma\delta_{CB}} + \frac{1-P}{\gamma+\delta_R+2\gamma\delta_{CB}}\right)}.$$

Substituting this expression into (B.3) and (B.4) for $j = L, R$, and after some algebra, we obtain the desired expressions. ■

Proof of Proposition 5

Direct computations yield $E\left(\pi^N\right) = \frac{P(m_R+2)A_L+(m_L+2)(1-P)A_R}{\Delta^N}$ and $E\left(\pi^D\right) = \frac{P(c_R m_R+2)A_L+(c_L m_L+2)(1-P)A_R}{\Delta^D}$. In addition, applying lemma A.1 for $z = \pi^N$ and $z = \pi^D$, it follows that

$$E\left(\left(\pi^N\right)^2\right) = \frac{\begin{pmatrix} P\left(\left(P + m_R + 1\right) A_L + \left(1 - P\right) A_R\right)^2 \\ +(1-P)((1 - P + m_L + 1)\, A_R + PA_L)^2 \end{pmatrix}}{\left(\Delta^N\right)^2} + \left(P\left(\frac{1}{m_L+2}\right)^2 + (1-P)\left(\frac{1}{m_R+2}\right)^2\right)\sigma_\varepsilon^2 \text{ and}$$

$$E\left((\pi^D)^2\right) = \frac{\begin{pmatrix} P\left(\left(c_R m_R + 1 + P\right) A_L + \left(1 - P\right) A_R\right)^2 \\ +(1-P)((c_L m_L + 2 - P)\, A_R + PA_L)^2 \end{pmatrix}}{\left(\Delta^D\right)^2} + \left(P\left(\frac{1}{c_L m_L+2}\right)^2 + (1-P)\left(\frac{1}{c_R m_R+2}\right)^2\right)\sigma_\varepsilon^2.$$

(a) Let $f(c_L, c_R) = E\left(\pi^D\right)$ and $g(c_L, c_R) = E\left((\pi^D)^2\right)$. Notice that $f(c_L, c_R)$ and $g(c_L, c_R)$ are decreasing functions in c_L and c_R. Moreover, $f(1,1) = E\left(\pi^N\right)$ and $g(1,1) = E\left((\pi^N)^2\right)$. The combination of these results allows us to conclude that $E\left(\pi^N\right) > E\left(\pi^D\right)$ and $E\left((\pi^N)^2\right) > E\left((\pi^D)^2\right)$ whenever $c_L > c_R \geq 1$.

(b) Suppose now that $c_L > 1 > c_R$. First, we focus on the comparison of the expected inflation. Let $h(P) = E\left(\pi^N\right) - E\left(\pi^D\right) = \frac{P(m_R+2)A_L+(m_L+2)(1-P)A_R}{\Delta^N} - \frac{P(c_R m_R+2)A_L+(c_L m_L+2)(1-P)A_R}{\Delta^D}$. Differentiating,

$$\frac{\partial}{\partial P}h(P) = \frac{(m_L+2)(m_R+2)(A_L(m_R+1) - A_R(m_L+1))}{(\Delta^N)^2} - \frac{\begin{pmatrix}(c_L m_L+2)(c_R m_R+2) \\ \times (A_L(c_R m_R+1) - A_R(c_L m_L+1))\end{pmatrix}}{(\Delta^D)^2}.$$

Now, we distinguish two cases:

Case 1: $A_L(c_R m_R+1) \leq A_R(c_L m_L+1)$. In this case $\frac{\partial}{\partial P}h(P) > 0$. Therefore, $h(P)$ is an increasing function in P. Moreover, $h(1) > 0$ and $h(0) < 0$ whenever $c_L > 1 > c_R$. This implies that there exists a unique value $\overline{P}$ such that $h(P) > 0$ (or equivalently, $E\left(\pi^N\right) > E\left(\pi^D\right)$) if and only if $P > \overline{P}$.

Case 2: $A_L(c_R m_R+1) > A_R(c_L m_L+1)$. In this case $\frac{\partial^2}{\partial^2 P}h(P) > 0$. Therefore, $h(P)$ is a convex function in P. Moreover, $h(1) > 0$ and $h(0) < 0$ whenever $c_L > 1 > c_R$. This implies that there exists a unique value $\overline{P}$ such that $h(P) > 0$ (or equivalently, $E\left(\pi^N\right) > E\left(\pi^D\right)$) if and only if $P > \overline{P}$.

In relation to the comparison of the term related to inflation stabilization, notice that as $c_L > 1 > c_R$, $E\left((\pi^N)^2\right)|_{P=1} > E\left((\pi^D)^2\right)|_{P=1}$ and $E\left((\pi^N)^2\right)|_{P=0} < E\left((\pi^D)^2\right)|_{P=0}$. Hence, we can conclude that when P is large enough, then $E\left((\pi^D)^2\right) < E\left((\pi^N)^2\right)$. ■

Proof of Lemma 4

(a) Substituting the expressions of π_L^N and π_R^N into the expressions of x_L^N and x_R^N, and taking expectations, we have $E\left(x_L^N\right) = x^* - \frac{(P+m_R+1)A_L+(1-P)A_R}{2\delta_L\Delta^N}$ and $E\left(x_R^N\right) = x^* - \frac{(1-P+m_L+1)A_R+PA_L}{2\delta_R\Delta^N}$. Using the expressions of m_L and m_R,

it follows that $E\left(x_L^N\right) > E\left(x_R^N\right)$ is equivalent to $\delta_L > \delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)}$.

(b) Substituting the expressions of π_L^D and π_R^D into the expressions of x_L^D and x_R^D and taking expectations, we have $E\left(x_L^D\right) = x^* - \frac{c_L}{2\delta_L}\frac{(c_R m_R+1+P)A_L+(1-P)A_R}{\Delta^D}$ and $E\left(x_R^D\right) = x^* - \frac{c_R}{2\delta_R}\frac{(c_L m_L+1+1-P)A_R+PA_L}{\Delta^D}$. From the expressions of m_L, m_R, c_L, and c_R, we have that $E(x_L^D) > E(x_R^D)$ if and only if $\delta_L > \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$. ■

Proof of Proposition 6

(a) Recall that $Var_P(x^N) > Var_P(x^D)$ is equivalent to $\left|E\left(x_L^N\right) - E\left(x_R^N\right)\right| > \left|E(x_L^D) - E(x_R^D)\right|$. Moreover, when monetary policy is delegated to a moderately conservative central bank, we have

$$E\left(x_L^N\right) > E\left(x_L^D\right) \text{ and} \tag{B.5}$$

$$E\left(x_R^N\right) < E\left(x_R^D\right). \tag{B.6}$$

Now we distinguish three cases: (1) $\delta_L \leq \delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)}$, (2) $\delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)} < \delta_L \leq \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$, and (3) $\delta_L > \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$.

Case 1: $\delta_L \leq \delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)}$. Using the proof of lemma 4, we have $E\left(x_L^N\right) \leq E\left(x_R^N\right)$. Combining this inequality, (B.5) and (B.6), it follows that $E\left(x_L^D\right) < E\left(x_L^N\right) \leq E\left(x_R^N\right) < E\left(x_R^D\right)$. Hence, $E\left(x_R^D\right) - E\left(x_L^D\right) > E\left(x_R^N\right) - E\left(x_L^N\right) \geq 0$, which implies that $Var_P(x^D) > Var_P(x^N)$.

Case 2: $\delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)} < \delta_L \leq \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$. Using the proof of lemma 4, we know

that $E\left(x_L^N\right) > E\left(x_R^N\right)$ and $E\left(x_L^D\right) \leq E\left(x_R^D\right)$. Therefore, to show $Var_P(x^N) < Var_P(x^D)$, it suffices to prove

$$E\left(x_L^N\right) - E\left(x_R^N\right) < E\left(x_R^D\right) - E\left(x_L^D\right). \tag{B.7}$$

Substituting the expressions of $E\left(x_L^N\right)$, $E\left(x_R^N\right)$, $E\left(x_R^D\right)$, and $E\left(x_L^D\right)$ in (B.7) and doing some algebra, we have that (B.7) is equivalent to $f(\delta_L) < g(\delta_L)$, where

$$f(\delta_L) = \frac{\begin{pmatrix} \gamma\left(\left(A_R + 4\gamma A_R + 2\gamma P\left(A_L - A_R\right)\right)\left(\delta_L - \delta_R\right) \\ \quad - \left(\gamma + \delta_R + 2\gamma\delta_R\right)\left(A_L - A_R\right)\right) \end{pmatrix}}{\begin{pmatrix} P\left(\gamma + \delta_R + 4\gamma\delta_R\right)\left(\gamma + \delta_L + 2\gamma\delta_L\right) \\ \quad + (1-P)\left(\gamma + \delta_R + 2\gamma\delta_R\right)\left(\gamma + \delta_L + 4\gamma\delta_L\right) \end{pmatrix}}$$

and

$$g(\delta_L) = \frac{-\gamma\left(A_R\left(\delta_L - \delta_R\right) - \left(A_L - A_R\right)\left(\gamma + \delta_R + \gamma\delta_{CB}\right)\right)}{\begin{pmatrix} P\left(\gamma + \delta_R + 2\gamma\delta_{CB}\right)\left(\gamma + \delta_L + \gamma\delta_{CB}\right) \\ \quad + (1-P)\left(\gamma + \delta_R + \gamma\delta_{CB}\right)\left(\gamma + \delta_L + 2\gamma\delta_{CB}\right) \end{pmatrix}}.$$

Direct computations yield that $f(\delta_L)$ is an increasing function in δ_L, while $g(\delta_L)$ is a decreasing function in δ_L. Furthermore,

$$\begin{aligned} & f\left(\delta_R + \frac{\left(\gamma + \delta_R + 2\gamma\delta_R\right)\left(A_L - A_R\right)}{\left(4\gamma + 1\right)A_R + 2P\gamma\left(A_L - A_R\right)}\right) \\ & \quad < g\left(\delta_R + \frac{\left(\gamma + \delta_R + 2\gamma\delta_R\right)\left(A_L - A_R\right)}{\left(4\gamma + 1\right)A_R + 2P\gamma\left(A_L - A_R\right)}\right) \text{ and} \\ & f\left(\delta_R + \frac{\left(\gamma + \delta_R + \gamma\delta_{CB}\right)\left(A_L - A_R\right)}{A_R}\right) \\ & \quad > g\left(\delta_R + \frac{\left(\gamma + \delta_R + \gamma\delta_{CB}\right)\left(A_L - A_R\right)}{A_R}\right). \end{aligned}$$

Hence, we conclude that there exists a unique value $\overline{\delta_L}$ belonging to the interval

$$\left(\delta_R + \frac{(\gamma + \delta_R + 2\gamma\delta_R)(A_L - A_R)}{(4\gamma + 1)A_R + 2P\gamma(A_L - A_R)}, \right.$$
$$\left. \delta_R + \frac{(\gamma + \delta_R + \gamma\delta_{CB})(A_L - A_R)}{A_R}\right)$$

such that $f(\delta_L) < g(\delta_L)$ if and only if $\delta_L \in \left(\delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{A_R(4\gamma+1)+2P\gamma(A_L-A_R)}, \overline{\delta_L}\right)$. Therefore, $Var_P(x^N) < Var_P(x^D)$ if and only if $\delta_L \in \left(\delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)}, \overline{\delta_L}\right)$.

Case 3: $\delta_L > \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$. In this case we know that $E\left(x_L^N\right) > E\left(x_R^N\right)$ and $E\left(x_L^D\right) > E\left(x_R^D\right)$. Therefore, in order to show $Var_P(x^N) > Var_P(x^D)$, it suffices to prove $E\left(x_L^N\right) - E\left(x_R^N\right) > E\left(x_L^D\right) - E\left(x_R^D\right)$ or, equivalently, $E\left(x_L^N\right) - E\left(x_L^D\right) > E\left(x_R^N\right) - E\left(x_R^D\right)$. Using (B.5) and (B.6), we know that the left-hand side of the previous inequality is positive and the right-hand side is negative. Hence, $Var_P(x^N) > Var_P(x^D)$.

(b) Consider now an ultraconservative central bank. Again, $Var_P(x^N) > Var_P(x^D)$ if and only if $\left|E\left(x_L^N\right) - E\left(x_R^N\right)\right| > \left|E(x_L^D) - E(x_R^D)\right|$. We distinguish three cases: (1) $\delta_L \leq \delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)}$, (2) $\delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{(4\gamma+1)A_R+2P\gamma(A_L-A_R)} < \delta_L \leq \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$, and (3) $\delta_L > \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$.

Case 1: $\delta_L \leq \delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{A_R(4\gamma+1)+2P\gamma(A_L-A_R)}$. In this case we know that $E\left(x_L^N\right) \leq E\left(x_R^N\right)$ and $E\left(x_L^D\right) < E\left(x_R^D\right)$. Therefore, in order to show $Var_P(x^N) < Var_P(x^D)$ it suffices to prove $E\left(x_R^N\right) - E\left(x_L^N\right) < E\left(x_R^D\right) - E\left(x_L^D\right)$ or, equivalently,

$$E\left(x_R^N\right) - E\left(x_R^D\right) < E\left(x_L^N\right) - E\left(x_L^D\right). \tag{B.8}$$

Notice that the right-hand side of the previous inequality is positive since it always holds $E\left(x_L^N\right) > E\left(x_L^D\right)$. Next, we distinguish two subcases: (1.1) $E\left(x_R^N\right) \leq E\left(x_R^D\right)$, and (1.2) $E\left(x_R^N\right) > E\left(x_R^D\right)$.

- *Subcase 1.1:* $\left(E\left(x_R^N\right) \leq E\left(x_R^D\right)\right)$. In this case (B.8) holds since the right-hand side of (B.8) is positive, whereas the left-hand side of (B.8) is negative. Consequently, $Var_P(x^N) < Var_P(x^D)$.
- *Subcase 1.2:* $\left(E\left(x_R^N\right) > E\left(x_R^D\right)\right)$. In this case, substituting the expressions of $E\left(x_L^N\right)$, $E\left(x_R^N\right)$, $E\left(x_R^D\right)$, and $E\left(x_L^D\right)$ and after some algebra, we have that (B.8) is equivalent to

$$\frac{\delta_L}{\delta_R} < \frac{(P + m_R + 1)\, A_L + (1 - P)\, A_R}{(1 - P + m_L + 1)\, A_R + PA_L} + \frac{\Delta^N F(c_L, c_R)}{\begin{pmatrix} (c_R \Delta^N((1 - P + c_L m_L + 1)A_R + PA_L) \\ \quad - \Delta^D((1 - P + m_L + 1)A_R + PA_L)) \\ \quad \times ((1 - P + m_L + 1)A_R + PA_L) \end{pmatrix}},$$

where

$$\begin{aligned} F(c_L, c_R) &= c_R(c_R - 1) \\ &\quad \times (P(A_L - A_R)(A_L m_R - A_R m_L) \\ &\quad + A_R(2A_L m_R - A_L m_L - A_R m_L)) + (c_L - c_R) \\ &\quad \times \begin{pmatrix} m_L(c_R - 1)(A_L - A_R)(A_R + P(A_L - A_R)) \\ \quad + (2A_R + P(A_L - A_R)) \\ \quad \times (A_L + A_R + P(A_L - A_R) + A_L m_L) \\ \quad + A_L c_R(m_R - m_L)(2A_R + P(A_L - A_R)) \end{pmatrix}. \end{aligned}$$

In this case the second term of the right-hand side of the previous inequality is positive. Then, we can conclude that this inequality is satisfied whenever $\frac{\delta_L}{\delta_R} \leq \frac{(P+m_R+1)A_L+(1-P)A_R}{(1-P+m_L+1)A_R+PA_L}$. Using the expressions of m_L and m_R, the previous inequality is equivalent to $\delta_L \leq \delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{A_R(4\gamma+1)+2P\gamma(A_L-A_R)}$. Consequently, we have that in this case $Var_P(x^N) < Var_P(x^D)$.

Case 2: $\delta_R + \frac{(\gamma+\delta_R+2\gamma\delta_R)(A_L-A_R)}{A_R(4\gamma+1)+2P\gamma(A_L-A_R)} < \delta_L \leq \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$. This proof is omitted since it is identical to the proof of case 2 in part (a).

Case 3: $\delta_L > \delta_R + \frac{(\gamma+\delta_R+\gamma\delta_{CB})(A_L-A_R)}{A_R}$. In this case we know that $E\left(x_L^N\right) > E\left(x_R^N\right)$ and $E\left(x_L^D\right) > E\left(x_R^D\right)$. Therefore, to show $Var_P(x^N) > Var_P(x^D)$, it suffices to prove $E\left(x_L^N\right) - E\left(x_R^N\right) > E\left(x_L^D\right) - E\left(x_R^D\right)$ or, equivalently, $E\left(x_L^N\right) - E\left(x_L^D\right) > E\left(x_R^N\right) - E\left(x_R^D\right)$. Substituting the expressions of $E\left(x_L^N\right)$, $E\left(x_R^N\right)$, $E\left(x_R^D\right)$, and $E\left(x_L^D\right)$ and after some algebra, we have that

$$\begin{aligned} & E\left(x_L^N\right) - E\left(x_L^D\right) - \left(E\left(x_R^N\right) - E\left(x_R^D\right)\right) \\ & = \frac{p(\delta_{CB})}{4\gamma^2\delta_L\delta_R\delta_{CB}^2\Delta^N\Delta^D}, \end{aligned}$$

with

$$\begin{aligned} & p(\delta_{CB}) \\ & = 2A_R(\delta_L - \delta_R)(P\gamma(\delta_L - \delta_R) \\ & \quad + (2\gamma^2 + \gamma\delta_R - \delta_L\delta_R(4\gamma+1))) \\ & \quad + 2\delta_L(A_L - A_R)(\gamma + \delta_R)(2\gamma + 2\delta_R(2\gamma+1) \\ & \quad + P(\delta_L - \delta_R))) \\ & \quad + \delta_{CB}\begin{pmatrix} \delta_L^2\left(A_R\left(4\gamma+1\right) + 2P\gamma\left(A_L - A_R\right)\right)\left(P+1\right) \\ + \delta_L\begin{pmatrix} A_R\left(4\gamma+1\right)\left(3\gamma + \delta_R - 2P\delta_R\right) + \\ \left(A_L - A_R\right) \\ \times\left(-4\gamma\delta_R P^2 + \left(4\gamma^2 - \gamma - \delta_R\right)P\right) \\ + 4\gamma\left(\gamma + \delta_R + 2\gamma\delta_R\right)) \end{pmatrix} \\ + \delta_R A_R\left(4\gamma+1\right)\left(P\delta_R - 2\delta_R - 3\gamma\right) \\ + \left(\gamma + \delta_R + 2\gamma\delta_R + 2P\gamma\delta_R\right) \\ \times\left(2\gamma + 2\delta_R - P\delta_R\right)\left(A_R - A_L\right) \end{pmatrix} \\ & \quad + \delta_{CB}^2\begin{pmatrix} 2\left(\delta_L - \left(\delta_R + \frac{(\gamma+\delta_R)(A_L-A_R)}{A_R}\right)\right) \\ \times\left(2\gamma P\left(A_L - A_R\right) + 4\gamma A_R + A_R\right)\gamma \\ + 4\dfrac{\begin{pmatrix} \left(\gamma P\left(A_L - A_R\right) + 2\gamma A_R + A_R\delta_R \right. \\ \left. + \delta_R P\left(A_L - A_R\right)\right)\left(A_L - A_R\right)\gamma^2 \end{pmatrix}}{A_R} \end{pmatrix}. \end{aligned}$$

Taking into account that $\delta_L > \delta_R + \frac{(\gamma+\delta_R)(A_L-A_R)}{A_R}$ and $A_L > A_R$, computations yield that the coefficients of δ_{CB}^2 and δ_{CB} are positive. In addition if $A_L - A_R$ is very close to 0, then we obtain the same results as in the benchmark case. Otherwise, i.e., if $A_L - A_R$ is high enough, then we obtain that the independent coefficient is positive. Thus, in this case we have that $p(\delta_{CB}) > 0$ and, hence, $Var_P(x^N) > Var_P(x^D)$. ■

Appendix C

Proposition C.1. *Under delegation in the sequential game, in equilibrium the inflation rates and outputs are given by*

$$\begin{aligned}
\pi_L^{D,S} &= \frac{\gamma\left(2\delta_{CB}+1\right)\delta_{CB}\left(\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2\right)}{\Delta^{D,S}}A \\
&\quad - \frac{\delta_{CB}\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\varepsilon, \\
\pi_R^{D,S} &= \frac{\gamma\left(2\delta_{CB}+1\right)\delta_{CB}\left(\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2\right)}{\Delta^{D,S}}A \\
&\quad - \frac{\delta_{CB}\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\varepsilon, \\
x_L^{D,S} &= x^* - \frac{1}{\delta_{CB}}\pi_L^{D,S} \text{ and } x_R^{D,S} = x^* - \frac{1}{\delta_{CB}}\pi_R^{D,S},
\end{aligned}$$

where

$$\begin{aligned}
\Delta^{D,S} &= P\left(\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2\right) \\
&\quad \times\left(\gamma+\delta_L+3\gamma\delta_{CB}+2\gamma\delta_{CB}^2+\delta_{CB}^2\right) \\
&\quad + (1-P)\left(\gamma+\delta_R+3\gamma\delta_{CB}+2\gamma\delta_{CB}^2+\delta_{CB}^2\right) \\
&\quad \times\left(\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2\right).
\end{aligned}$$

Proof of Proposition C.1. We solve by backward induction. In the second step, the central bank chooses the inflation rate after observing the tax rate chosen by the government in office. Suppose that

party j is in office. Under delegation, the central bank chooses π_j in order to solve the following problem after observing τ_j:

$$\underset{\pi_j}{min} V_{CB} = \frac{1}{2}\left(\pi_j^2 + \delta_{CB}\left(x_j - x^*\right)^2\right).$$

As in the simultaneous setup, the FOC is given by $\frac{\partial}{\partial \pi_j} V_{CB} = \pi_j + \delta_{CB}\left(x_j - x^*\right) = 0$, which implies

$$\pi_j = \frac{\delta_{CB}}{\delta_{CB}+1}\left(w^* + x^* + \pi^e + \tau_j - \varepsilon\right). \tag{C.1}$$

In the first step, the fiscal authority chooses τ_j taking into account (C.1). Thus,

$$\begin{aligned} \underset{\tau_j}{min} V_{G_j} &= \tfrac{1}{2}\left(\pi_j^2 + \delta_j\left(x_j - x^*\right)^2 + \gamma(g_j - g^*)^2\right) \\ \text{s.t.} \quad \pi_j &= \tfrac{\delta_{CB}}{\delta_{CB}+1}\left(w^* + x^* + \pi^e + \tau_j - \varepsilon\right). \end{aligned}$$

The FOC of this problem is given by

$$\begin{aligned} \frac{\partial}{\partial \tau_j} V_{G_j} = \pi_j \frac{\delta_{CB}}{\delta_{CB}+1} + \delta_j\left(x_j - x^*\right)\left(\frac{\delta_{CB}}{\delta_{CB}+1} - 1\right) \\ + \gamma(g_j - g^*)\left(1 + \frac{\delta_{CB}}{\delta_{CB}+1}\right) = 0. \end{aligned}$$

Using expressions (C.1), (1), and (2), it follows that

$$\tau_j = g^* - \frac{\delta_j + \gamma\delta_{CB} + 2\gamma\delta_{CB}^2 + \delta_{CB}^2}{\gamma + \delta_j + 4\gamma\delta_{CB} + 4\gamma\delta_{CB}^2 + \delta_{CB}^2}\left(A + \pi^e - \varepsilon\right).$$

Substituting the previous expression in (C.1),

$$\pi_j = \frac{\delta_{CB}\gamma\left(2\delta_{CB}+1\right)}{\gamma + \delta_j + 4\gamma\delta_{CB} + 4\gamma\delta_{CB}^2 + \delta_{CB}^2}\left(A + \pi^e - \varepsilon\right). \tag{C.2}$$

Using the previous formula in the expression for π^e and solving for π^e, we get

$$\pi^e = \frac{\left(\begin{array}{c}\gamma\left(2\delta_{CB}+1\right)\delta_{CB}\left(\frac{P}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right.\\ \left.+\frac{1-P}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right)\end{array}\right)}{\left(\begin{array}{c}1-\gamma\left(2\delta_{CB}+1\right)\delta_{CB}\left(\frac{P}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right.\\ \left.+\frac{1-P}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right)\end{array}\right)}A. \quad \text{(C.3)}$$

Substituting this expression into (C.2), for $j = L, R$, and after some algebra we obtain the desired expressions of the inflation rates. Finally, the expressions of outputs follow from the FOC of the optimization problem of the monetary authority. ■

Proof of Lemma 5

Using the expressions of $x_L^{D,S}$ and $x_R^{D,S}$ given in proposition C.1 and taking expectations, it follows that $E\left(x_L^{D,S}\right) - E\left(x_R^{D,S}\right) = \frac{\gamma(\delta_L-\delta_R)(2\delta_{CB}+1)}{\Delta^{D,S}}A$. Hence, we can conclude that $E\left(x_L^{D,S}\right) > E\left(x_R^{D,S}\right)$ since $\delta_L > \delta_R$. ■

Proof of Proposition 7

Let $h^S(P) = E\left(\pi^N\right) - E\left(\pi^{D,S}\right)$ and $f^S(P) = E\left((\pi^N)^2\right) - E\left((\pi^{D,S})^2\right)$. Using (A.3) and (C.3), it follows that $\frac{\partial}{\partial P}\left(E\left(\pi^N\right)\right) > 0$ and $\frac{\partial}{\partial P}\left(E\left(\pi^{D,S}\right)\right) < 0$. Therefore, $h^S(P)$ is an increasing function in P. Applying lemma A.1, we have

$$E\left((\pi^N)^2\right) = \left(\frac{\frac{P}{\frac{\frac{1}{\delta_L}+\frac{1}{\gamma}}{2}+2} + \frac{1-P}{\frac{\frac{1}{\delta_R}+\frac{1}{\gamma}}{2}+2}}{1-\left(\frac{P}{\frac{\frac{1}{\delta_L}+\frac{1}{\gamma}}{2}+2} + \frac{1-P}{\frac{\frac{1}{\delta_R}+\frac{1}{\gamma}}{2}+2}\right)}A\right)^2$$
$$+\left(P\left(\frac{1}{\frac{\frac{1}{\delta_L}+\frac{1}{\gamma}}{2}+2}\right)^2 + (1-P)\left(\frac{1}{\frac{\frac{1}{\delta_R}+\frac{1}{\gamma}}{2}+2}\right)^2\right)\sigma_\varepsilon^2$$

and

$$
\begin{aligned}
&E\left((\pi^{D,S})^2\right) \\
&= \left(\frac{\left(\begin{array}{c} \gamma\left(2\delta_{CB}+1\right)\delta_{CB}\left(\frac{P}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right) \\ +\frac{1-P}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\end{array}\right)}{\left(\begin{array}{c} 1-\gamma\left(2\delta_{CB}+1\right)\delta_{CB}\left(\frac{P}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right) \\ +\frac{1-P}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\end{array}\right)} A \right)^2 \\
&+ \left(P\left(\frac{\delta_{CB}\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right)^2 \right. \\
&\left. + (1-P)\left(\frac{\delta_{CB}\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right)^2 \right)\sigma_\varepsilon^2.
\end{aligned}
$$

Direct computations yield that $\frac{\partial}{\partial P}\left(E\left((\pi^N)^2\right)\right) > 0$ and $\frac{\partial}{\partial P}\left(E\left((\pi^{D,S})^2\right)\right) < 0$. Therefore, $f^S(P)$ is an increasing function in P. Next, we distinguish two cases:

(a) The central bank is ultraconservative ($\delta_{CB} \leq 2\delta_R$). Combining the monotonicity property of $h^S(P)$ and $f^S(P)$ and the fact that $h^S(0) \geq 0$ and $f^S(0) \geq 0$ when $\delta_{CB} \leq 2\delta_R$, we have that $h^S(P) > 0$ and $f^S(P) > 0$ or, equivalently, $E\left(\pi^N\right) > E\left(\pi^{D,S}\right)$ and $E\left((\pi^N)^2\right) > E\left((\pi^{D,S})^2\right)$, whenever $P > 0$.

(b) The central bank is moderately conservative ($2\delta_R < \delta_{CB} < 2\delta_L$). Combining the monotonicity property of $h^S(P)$ and the fact that $h^S(0) < 0$, $f^S(0) < 0$, $h^S(1) > 0$, and $f^S(1) > 0$ whenever $2\delta_R < \delta_{CB} < 2\delta_L$, it follows that there exists a unique value $\overline{P^S}$ such that $h^S(P) > 0$ and $f^S(P) > 0$ or, equivalently, $E\left(\pi^N\right) > E\left(\pi^{D,S}\right)$ and $E\left((\pi^N)^2\right) > E\left((\pi^{D,S})^2\right)$ whenever $P > \overline{P^S}$. ■

Proof of Proposition 8

(a) Recall that $Var_P(x^N) > Var_P(x^{D,S})$ if and only if $E\left(x_L^N\right) - E\left(x_R^N\right) > E\left(x_L^{D,S}\right) - E\left(x_R^{D,S}\right)$. Let $g^S(\delta_{CB}) = E\left(x_L^N\right) - E\left(x_R^N\right) - \left(E\left(x_L^{D,S}\right) - E\left(x_R^{D,S}\right)\right)$. Direct computations yield that g^S is increasing in δ_{CB}. Next, we distinguish two cases:

Case 1: The central bank is moderately conservative ($2\delta_R < \delta_{CB} < 2\delta_L$). Combining the monotonicity property of $g^S(\delta_{CB})$ and the fact that $g^S(2\delta_R) > 0$, we have that $g^S(\delta_{CB}) > 0$ whenever $2\delta_R < \delta_{CB} < 2\delta_L$ and, hence, $Var_P(x^N) > Var_P(x^{D,S})$.

Case 2: The central bank is ultraconservative ($\delta_{CB} \leq 2\delta_R$). Then, we consider two subcases:

- *Subcase 2.1:* $g^S(0) \geq 0$. In this case $2\gamma^2 + \gamma\left(\delta_R + P\delta_L - P\delta_R - 4\delta_L\delta_R\right) - \delta_L\delta_R \geq 0$. Descartes' rule tells us that there exists a unique value of γ, denoted by $\overline{\gamma^S}$, such that the previous inequality is satisfied whenever $\gamma \geq \overline{\gamma^S}$.[24] Combining the monotonicity property of $g^S(\delta_{CB})$ and the fact that $g^S(0) \geq 0$, we have that $g^S(\delta_{CB}) > 0$ whenever $\delta_{CB} > 0$ and, hence, $Var_P(x^N) > Var_P(x^{D,S})$. Therefore, we conclude that if $\gamma \geq \overline{\gamma^S}$, then $Var_P(x^N) > Var_P(x^{D,S})$.
- *Subcase 2.2:* $g^S(0) < 0$ (or, equivalently, $\gamma < \overline{\gamma^S}$). In this case, the monotonicity property of $g^S(\delta_{CB})$ implies that there exists a unique value of δ_{CB}, denoted by $\overline{\delta_{CB}^S}$, such that $g^S(\delta_{CB}) > 0$ whenever $\delta_{CB} > \overline{\delta_{CB}^S}$. Consequently, $Var_P(x^N) > Var_P(x^{D,S})$ if and only if $\delta_{CB} > \overline{\delta_{CB}^S}$.

(b) From the proof of proposition C.1, it follows that

$$x_L^{D,S} = x^* - \frac{1}{\delta_{CB}} \frac{\begin{pmatrix} \gamma\left(2\delta_{CB}+1\right)\delta_{CB} \\ \times\left(\gamma + \delta_R + 4\gamma\delta_{CB} + 4\gamma\delta_{CB}^2 + \delta_{CB}^2\right) \end{pmatrix}}{\Delta^{D,S}} A$$

[24] Note that $\overline{\gamma^S}$ coincides with $\overline{\gamma}$ given in the proof of proposition 3.

$$+\frac{\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\varepsilon \text{ and}$$

$$x_R^{D,S}=x^*-\frac{1}{\delta_{CB}}\frac{\begin{pmatrix}\gamma\left(2\delta_{CB}+1\right)\delta_{CB}\\ \times\left(\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2\right)\end{pmatrix}}{\Delta^{D,S}}A$$

$$+\frac{\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\varepsilon.$$

Applying lemma A.1, we have that

$$\begin{aligned}&Var_E\left(x^{D,S}\right)\\&=\left(P\left(\frac{\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_L+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right)^2\right.\\&\quad\left.+\left(1-P\right)\left(\frac{\gamma\left(2\delta_{CB}+1\right)}{\gamma+\delta_R+4\gamma\delta_{CB}+4\gamma\delta_{CB}^2+\delta_{CB}^2}\right)^2\right)\sigma_\varepsilon^2.\end{aligned}$$

In addition, in the proof of proposition 4 we have obtained

$$\begin{aligned}&Var_E\left(x^N\right)\\&=\left(P\left(\frac{1}{2\delta_L\left(m_L+2\right)}\right)^2+\left(1-P\right)\left(\frac{1}{2\delta_R\left(m_R+2\right)}\right)^2\right)\sigma_\varepsilon^2\end{aligned}$$

or, using the expressions of m_L and m_R,

$$\begin{aligned}&Var_E\left(x^N\right)\\&=\left(P\left(\frac{\gamma}{\gamma+\delta_L+4\gamma\delta_L}\right)^2+\left(1-P\right)\left(\frac{\gamma}{\gamma+\delta_R+4\gamma\delta_R}\right)^2\right)\sigma_\varepsilon^2.\end{aligned}$$

Hence, $Var_E\left(x^N\right)-Var_E\left(x^D\right)$ is a linear function in P. Next we distinguish two cases:

Case 1: $c_R<1<c_L$ (i.e., $2\delta_R<\delta_{CB}<2\delta_L$). Direct computations yield that $Var_E\left(x^N\right)|_{P=1}-Var_E\left(x^D\right)|_{P=1}<0$ and that $Var_E\left(x^N\right)|_{P=0}-Var_E\left(x^D\right)|_{P=0}>0$ whenever $2\delta_R<\delta_{CB}<2\delta_L$. Hence, we can conclude that there exists

a unique value of P, denoted by $\overline{\overline{P^S}}$, such that $Var_E\left(x^D\right) > Var_E\left(x^N\right)$ if and only if $P > \overline{\overline{P^S}}$.

Case 2: $c_L > c_R \geq 1$ (i.e., $\delta_{CB} \leq 2\delta_R < 2\delta_L$). Note that $Var_E\left(x^N\right)|_{P=1} - Var_E\left(x^D\right)|_{P=1} < 0$ and that $Var_E\left(x^N\right)|_{P=0} - Var_E\left(x^D\right)|_{P=0} \leq 0$ whenever $\delta_{CB} \leq 2\delta_R < 2\delta_L$. Therefore, for all $0 < P < 1$, it holds that $Var_E\left(x^D\right) > Var_E\left(x^N\right)$ whenever $\delta_{CB} \leq 2\delta_R < 2\delta_L$. ■

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Residential Investment and Economic Activity: Evidence from the Past Five Decades*

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We analyze the evolution and main drivers of residential investment in 15 advanced economies using a large panel with quarterly data since the 1970s. Residential investment is a notably volatile component of real GDP in all countries in the sample. Real house price growth, net migration inflows, household size, and the existing housing stock are significant drivers of residential investment across various model specifications. We detect important asymmetries: interest rate increases affect residential investment more than interest rate declines, and interest rate changes have larger effects on residential investment when its share in GDP is rising. We also show that information on residential investment significantly improves the performance of standard recession-prediction models.

JEL Codes: E22, E32, E37, E43, E52, F44.

1. Introduction

Most research on residential investment focuses on how booms and busts in house prices and housing credit affect macroeconomic and financial outcomes. For example, Jordà, Schularick, and

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Taylor (2016) documented how mortgage credit booms in advanced economies since the Second World War were increasingly associated with deeper recessions and slower recoveries. The co-movement of property prices and credit also features prominently in financial cycle models, whose peaks are closely associated with financial crises (Claessens, Kose, and Terrones 2012; Borio 2014).

Research on the macroeconomic implications of residential investment in terms of volumes—housing output produced for the market—is much scarcer. This is surprising given that residential investment is one of the most volatile components of gross domestic product (GDP), and given the housing booms and busts in countries such as the United States, Ireland, and Spain in the 2000s, or the Nordic countries in the 1990s. There are also very few cross-country analyses on the drivers of residential investment, including the role of interest rates and their potentially asymmetric effects over the cycle. For policymakers this is an important issue: for instance, in an economy burdened by oversupply of housing after a real estate boom, lower interest rates may do little to kick-start residential investment and economic activity.

This paper intends to fill part of this gap. One contribution we make is to examine the proximate drivers of residential investment in a cross-country rather than single-country context. Another is that we conduct the analysis on data over the past half-century. This helps us uncover common financial, demographic, and real economy factors related to both demand and supply, rather than idiosyncratic, country- or time-specific factors. We also study the behavior of residential investment across the business cycle, in particular the leading indicator properties of residential investment in a simple benchmark recession-prediction model.

The paper highlights three main findings. First, we show that the main determinants of residential investment in advanced economies are real house prices, nominal interest rates, demographic factors, and the state of housing supply. House prices seem to play a prominent role by affecting the incentives to pursue residential investment, i.e., the numerator in Tobin's q for housing.

Second, we find that the effects of interest rates on residential investment are twice as large during housing booms compared with normal times, and are clearly stronger when interest rates are rising than when they are falling.

Third, drops in residential investment consistently lead economic downturns in the 99 recessions identified in our sample. This signaling property arises despite the small overall share of residential investment in GDP—around 6 percent on average over the past five decades. Prior to an economic downturn, house prices, growth in construction activity, and construction employment all decline. We show that information on declines in residential investment improves the performance of standard recession forecasting models that also feature the slope of the yield curve.

Our work is related to several strands of literature. One covers studies of residential investment over the business cycle—in particular, its recession prediction properties (Leamer 2007, 2015 and International Monetary Fund 2008). For instance, Leamer (2015) showed how 9 out of 11 recessions in the United States after the Second World War were preceded by large declines in residential investment. We confirm these properties in a broad cross-country setting, and highlight the economic mechanisms that may lie behind them. The relevance of housing dynamics for economic activity is consistent with the models of Iacoviello (2005) and Iacoviello and Neri (2010), in which housing wealth affects spending through changes in collateral values and hence borrowing constraints. More recently, Mian, Sufi, and Verner (2017) found that higher house price and household debt growth both predicted lower subsequent GDP growth. Herstad (2016) and Huang et al. (2018) also concluded that housing market developments predict business cycle variations. In addition, we build on the recession prediction models of Estrella and Hardouvelis (1991) and Rudebush and Williams (2009), who established that shifts in the yield curve anticipate downturns.

Another strand of the literature comprises studies of the determinants of residential investment. These have so far mainly focused on the United States. Our finding that residential investment in a broad range of advanced economies responds first and foremost to house price developments is in line with theoretical model predictions and empirical studies for individual countries (e.g., Topel and Rosen 1988, Tsoukis and Westaway 1994, Davis and Heathcote 2005, Glaeser, Gyourko, and Saiz 2008). Our study also relates to a number of papers that highlight the interest rate sensitivity of residential investment (e.g., McCarthy and Peach 2002, Erceg and Levin 2006, Jarocinski and Smets 2008, Aspachs-Bracons and

Rabanal 2011, Dokko et al. 2011, Calza, Monacelli, and Stracca 2013). One implication of this finding is that tighter monetary policy could have curtailed the magnitude of the pre-crisis housing boom in the United States (Taylor 2007, Leamer 2015, and Sutton, Mihaljek, and Subelyte 2017).

A third strand of the literature is studies of the asymmetric effects of interest rate changes on aggregate output (e.g., Angrist, Jordà, and Kuersteiner 2013, Vavra 2014, Tenreyro and Thwaites 2016). We find that the effects of interest rates on residential investment are stronger during housing booms and when interest rates are rising. We suggest that one source of this asymmetry could be downward house price rigidity, which forces adjustments through quantities rather than prices. Another is adjustment costs: labor shortages and other bottlenecks constrain the expansion of residential investment during the housing booms, but there are no constraints to slowing the activity during busts. More generally, our results corroborate the notion that booms that lead to a temporary oversupply of housing tend to be followed by periods of weak or unresponsive residential investment (Rognlie, Shleifer, and Simsek 2018). One implication of this feature of housing supply is that prolonged construction booms fueled by expansionary monetary policy could over time weaken the responsiveness of residential investment to monetary policy.

The paper is structured as follows. Section 2 describes the data and key stylized facts about residential investment. Section 3 outlines the empirical approach. Section 4 discusses the estimation results for determinants of residential investment. Section 5 presents results from a formal recession-prediction model that incorporates residential investment. Section 6 presents robustness tests. Section 7 concludes.

2. Data and Stylized Facts

The bulk of our residential investment data come from the Organisation for Economic Co-operation and Development's (OECD's) Economic Outlook database and national statistical authorities.[1]

[1]For Germany, pre-1991 data are for West Germany. For Switzerland, we use gross private domestic investment in construction.

To compute the size of the existing housing stock, we start from the estimates on initial housing stocks by Piketty and Zucman (2014). We then compute the stocks in subsequent periods by assuming a 1½ percent annual depreciation rate.[2]

For real house prices, we use residential property prices published by the Bank for International Settlements and deflated by the respective national consumer price indexes. For most countries in the sample, these series are based on transactions data. They also feature a quality adjustment based either on size or a more sophisticated approach such as hedonic regression (see Scatigna, Szemere, and Tsatsaronis 2014).

Our sample runs from 1970:Q1 to 2017:Q2 and is thus longer than in previous cross-country studies (e.g., International Monetary Fund 2008 and Calza, Monacelli, and Stracca 2013). We include 15 advanced economies: Australia, Canada, France, Germany, Italy, Japan, the Netherlands, New Zealand, Norway, Korea, Spain, Sweden, Switzerland, the United Kingdom, and the United States.

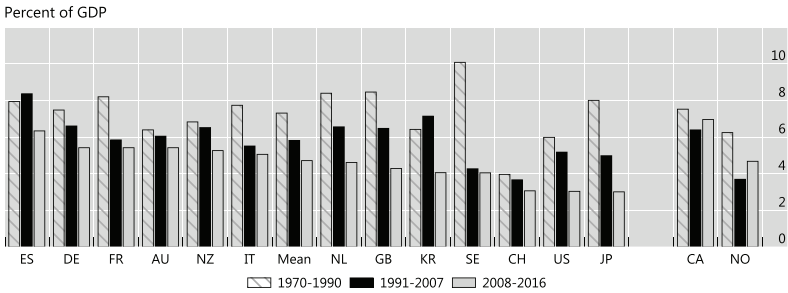
Residential investment generally accounts for a small share of GDP. Over the whole sample, it averaged 5.9 percent of GDP, compared with 12.5 percent for business investment. The share has fallen over time, from 7.3 percent in the first two decades to 5.8 percent between 1991 and the Great Financial Crisis (GFC), and then to 4.7 percent in 2008–16 (figure 1). Only Canada and Norway—shown on the right-hand side of the figure—have seen an increase in the share of residential investment post-crisis.

The share of residential investment in GDP was highest during the 1970s and 1980s, especially in Sweden (10 percent), where construction surged during the "Million Homes Programme" (1965–75) that aimed to overcome urbanization-induced housing shortages (Emanuelsson 2015). In the two decades before the GFC, the share of residential investment was the highest in Spain (8 percent), where easy financing conditions that accompanied the introduction of the euro, demographic factors, and purchases by other EU residents played a role (e.g., Garcia-Herrero and Fernandez de Lis 2008).

At the sectoral level, residential investment relates most closely to construction, whose value added in GDP has also generally fallen

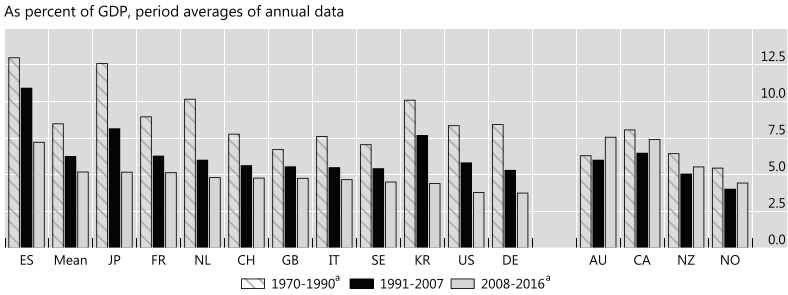
[2]In a robustness check we also estimated the model using a 2 percent depreciation rate. This change did not affect the results in a significant way.

Figure 1. Residential Investment over Time[a]



Sources: OECD, Economic Outlook database; national data; authors' calculations.
[a]Period averages of quarterly data. For Switzerland, annual data. Countries are ordered by the residential investment share in 2008–16.

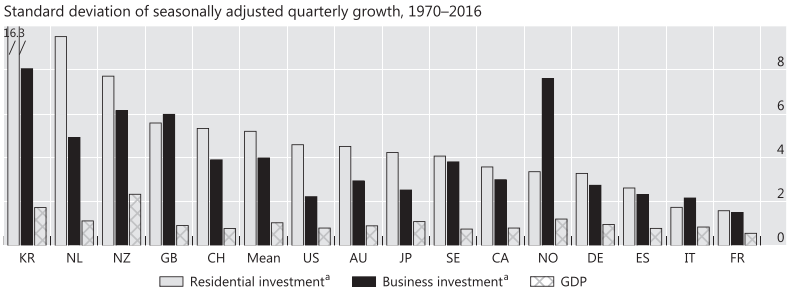
Figure 2. Construction Value Added



Sources: OECD; Datastream; national data; authors' calculations.
[a]For Japan, data are available to 2015; for New Zealand, data are available from 1971 to 2015. Countries are ordered by the construction value added share in 2008–16.

over time. In our sample, it averaged 8.5 percent in 1970–90; 6.2 percent in 1991–2007, and 5.2 percent in 2008–16 (figure 2). Post-crisis, the construction share of GDP has increased in Australia, Canada, New Zealand, and Norway, all of which have also seen housing booms during the period. Reflecting the high labor intensity of construction, the employment share has been relatively stable: it amounted on average to 7.8 percent of total employment in 1970–90 and 7.2 percent in 2008–16 (appendix figure A.1).

Figure 3. Residential and Business Investment Are Much More Volatile than GDP Growth[a]



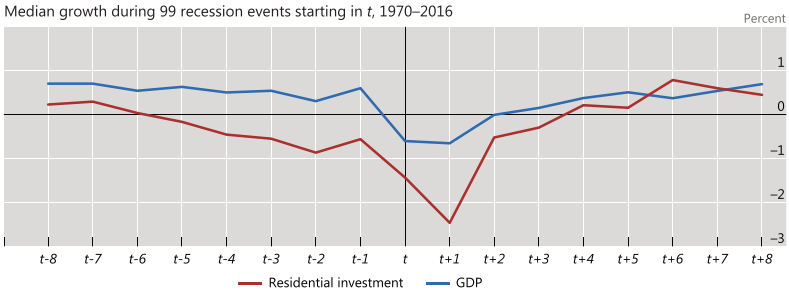
Sources: OECD, Economic Outlook database; national data; authors' calculations.
[a]Business investment includes investment in commercial property. For Germany (prior to 1991), Italy, and Spain, computed as total less housing investment. For Switzerland, construction is used for residential, and equipment and software for business investment.

Notwithstanding its generally small share in the overall economy, residential investment is often the most volatile component of GDP (figure 3). Volatility arises in part from the large housing stock: even small changes in desired housing stock require relatively large changes in investment. Measured by the standard deviation, volatility of residential investment in our sample is on average about five times that of overall GDP growth. This compares with a ratio of around four-to-one for business investment. Residential investment volatility in our sample has been highest in Korea and the Netherlands, at around eight to nine times that of GDP. Norway is an exception to this pattern, as growth in business investment, which is driven by oil production, is more volatile than that in residential investment. Yet, both series are far more volatile than GDP growth.

Consistent with high volatility, residential investment tends to fall sharply before recessions, with a notable lead over broader economic activity. Defining recessions conventionally, as a minimum of two consecutive periods of negative quarter-on-quarter GDP growth, figure 4 shows the median growth rates in real GDP (blue line) and residential investment (red line), before and after the start of a recession (in quarter t).[3] The sample includes 99 recessions. Residential

[3]For color versions of the figures, see the online version of the paper at http://www.ijcb.org.

Figure 4. Residential Investment versus GDP Growth during Recessions



Sources: OECD, Economic Outlook database; national data; authors' calculations.
Notes: Recession events are defined as at least two consecutive quarters of negative GDP growth based on seasonally adjusted data. t denotes the first quarter of recessions.

investment growth tends to become negative five quarters before the start of a recession, and then falls quite sharply before turning positive again four quarters after the start of a recession. Cumulatively, residential investment declines by a median of 7.3 percentage points of GDP during the quarters around a recession in our sample, while output falls by only 1.3 percentage points.[4]

Such dynamics are not merely mechanical, as real GDP declines over and above the fall in residential investment. We show this in appendix figure A.2, where the GDP series excludes the residential investment component. Moreover, the finding is robust to the inclusion of both country and time fixed effects, as shown in appendix figure A.3. These variables capture differences in the average growth rates between countries and control for any common shocks that may simultaneously affect residential investment and real GDP growth in several countries.

3. Empirical Methodology

To identify the main drivers of residential investment, we rely on direct single-equation estimations at multiple horizons. This

[4]While figure 4 shows growth outcomes during a median recession episode, the large cross-country data set features a rich set of recession events with diverse growth outcomes around such periods.

approach is flexible in that, contrary to a standard vector autoregression, it allows modeling asymmetric and nonlinear responses, which may occur due to pronounced booms and busts in the sample (see Agnello and Schuknecht 2011 and Dokko et al. 2011). This condition is fulfilled in our case given the long time span and wide geographic coverage of our data. More generally, this approach tends to be more robust to structural changes that may have occurred over the sample period.

Following the approach proposed by Jordà (2005), our estimated equations take the form

$$I_{i,t+h+1} - I_{i,t+h} = \alpha_i + \gamma_t + \beta' X_{i,t} + \varepsilon_{i,t+h+1}. \tag{1}$$

The equation is estimated by ordinary least squares, with inference based on standard errors that are clustered by country. The dependent variable is the change in the log of real residential investment I in country i between periods $t+h$ and $t+h+1$. We consider quarterly horizons from $h = 0$ to $h = 5$. The right-hand-side variables include country fixed effects, α_i, to capture country-specific unobserved heterogeneities; quarterly time fixed effects, γ_t, to capture common global trends; and a vector of explanatory variables, $X_{i,t}$. The explanatory variables are expressed in differences to avoid issues with nonstationarity and spurious correlations, and are lagged relative to the dependent variable to minimize reverse causality concerns. At the same time, we acknowledge that simultaneity issues cannot be fully eliminated in this estimating framework.[5]

The vector of explanatory variables $X_{i,t}$ includes various components of Tobin's q for residential investment that capture the incentives to invest in housing. Tobin's q for housing compares the market price of a home in the numerator with various costs of building a new home in the denominator. As in Tsoukis and Westaway (1994), we include the proxies for Tobin's q in the regression individually,

[5]It is important to note that including a lagged dependent variable did not lead to meaningful changes in the coefficients of other explanatory variables. This was found to be true irrespective of whether we used one or three lags of the dependent variable. Also, whether the housing stock variable (which is obviously correlated with the lagged dependent variable) was included or dropped did not produce material changes to the estimates. The lags of the dependent variable were generally not statistically significant and had coefficients below 0.10. These results are available upon request.

as there is considerable uncertainty regarding the measurement and weights of such components in any constructed measure of q.

Our house price series are computed as residential property prices deflated by the consumer price index. Time-to-build considerations imply that expected rather than observed house prices matter for investment decisions (Kydland and Prescott 1982). We estimate expected house prices from an adaptive expectations model: in the baseline regression we include house prices predicted by a simple AR(5) model, using lags of five quarters, based on the Akaike and Bayesian information criteria. As an alternative we simply use the last observation of real house price variation (results are available upon request). In theory, house price changes would be irrelevant for residential investment only if they were perceived to be entirely temporary. In practice, they are likely to carry new information about the trend and hence about rising or falling returns in the housing sector. The two series for expected house prices led to broadly similar results. Sutton, Mihaljek, and Subelyte (2017) argue that forecastable upward moves in house prices tend to persist because of the large search-and-transaction costs associated with house purchases. Similarly, Glaeser and Gyourko (2007) regard the serial correlation of house price changes as one of the key stylized facts of the housing market. In other words, the most recent changes contain information on the likely future trend in prices.

House prices should also reflect any additional effects of residential investment that are not well captured by other explanatory variables, and for which little comparable cross-country data are available. These include housing quality adjustments, the effects of restrictive regulations on housing supply, or, in some countries, the effects of foreign demand on residential investment. House prices also matter for the incentives to construct new housing and undertake home improvements, both of which are included in the residential investment series.

The components that enter the denominator of Tobin's q include three proxies for costs of residential investment. First, the producer price index (PPI) is a proxy for construction costs. Higher costs of raw materials, including energy, push up construction costs and could therefore negatively affect investment in new or the expansion of existing residential units. Second are short-term interest rates, which affect both housing supply through property developers'

funding costs, and housing demand through debt servicing costs. Interest rates also affect investment indirectly, through discount rates used in property valuations (the numerator of q). We use a short-term money market rate, in line with Mishkin's (2007) argument that builders construct houses relatively quickly, and express it in nominal terms, as money illusion phenomena may be important (Topel and Rosen 1988; Tsoukis and Westaway 1994).[6] Third, we include expected consumer price inflation, estimated by an adaptive AR(5) model, to allow for the possibility that rising inflation may depress investment, among other reasons because of the associated rise in uncertainty or macroeconomic instability.[7]

Among demand-side variables, we include income levels (GDP per capita), household size (number of persons per household), and net migration rates (net migrants per thousand residents). Higher income and greater inward migration are expected to boost residential investment (Monnet and Wolf 2017). Higher migration rates can also boost housing supply by increasing the number of construction workers.[8] Household size has a priori an ambiguous effect on residential investment: on the one hand, larger households normally imply higher density of living space, which should provide an incentive for developers to build more housing. On the other hand, smaller household size could reflect a cultural shift and over time might also boost demand. As an additional demographic variable we include the share of people aged 20 to 34 in total population, as the young are more likely to demand new housing units.

Finally, we include the value of the stock of housing in relation to GDP. A negative coefficient for this variable would imply that it effectively acts like a cointegration vector that pushes the economy back to its long-term equilibrium.[9] Larger existing stocks relative to the size of the economy should reduce the incentives to pursue additional residential investment.

[6] With real interest rates we also generally obtain the expected negative signs but with lower significance.

[7] Another variable that would enter the denominator of Tobin's q for housing is the price of land, as this affects the cost of producing a new home (see Corder 2008). Due to lack of comparable cross-country data on land prices, we did not include this variable in the estimation.

[8] Corder (2008) argued that labor shortages in the United Kingdom's construction sector eased in the 2000s when migrant workers entered the sector.

[9] We thank an anonymous referee for noting this point.

4. Estimation Results

4.1 Baseline Specifications

Table 1 presents our baseline estimates, based on around 2,700 country-quarter observations. t-statistics reported below the coefficients are based on robust standard errors clustered by country. Column 1 reports the drivers of residential investment growth between t and $t+1$, column 2 between $t+1$ and $t+2$, and so on. They indicate that several of our explanatory variables have an economically and statistically significant impact on residential investment over multiple horizons. This suggests that a sufficiently flexible modeling approach that considers a number of different lags is indeed appropriate. Furthermore, all coefficients have the expected sign whenever they are statistically significant at conventional levels.

Four findings are worth highlighting. First, higher real house prices are positively correlated with residential investment. A 1 percent increase in expected real house prices is associated with a 0.54 percent rise in residential investment already after one quarter, and 0.81 percent after two quarters. Changes in interest rates are negatively related to residential investment, with a lag. An increase in nominal interest rates by 100 basis points is associated with a 0.18 percent fall in residential investment after three quarters, and 0.37 percent after five quarters (if one considers coefficients that are at least significant at a 10 percent level).[10] Comparing normalized estimates instead (not shown), the relationship between residential investment and a one-standard-deviation change in real house prices is roughly twice as large, in absolute terms, as that of one-standard-deviation change in interest rates.[11]

Second, demographic factors are important. A higher rate of net migration is associated with more residential investment. Similarly,

[10]This effect is a bit smaller than the −0.6 percent reported in Calza, Monacelli, and Stracca (2013), who used a vector autoregressive (VAR) approach and data for 19 advanced economies.

[11]It is important to acknowledge that here we do not estimate the local supply elasticity separately. In countries in which the housing supply is inelastic, positive shocks to housing demand will increase prices but construction will rise by less. The degree of supply elasticity is also likely to affect the estimated relationship between lagged house prices and construction. We are indebted to an anonymous referee for noting this point.

Table 1. Determinants of Residential Investment Growth: 1970–2017

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.049	−0.037	−0.024	0.002	0.016	0.009
t-stat	1.06	1.09	0.89	0.07	0.58	0.42
Real House Price Growth Exp.	0.537***	0.276***	0.116	0.046	−0.046	0.009
t-stat	5.32	2.90	1.31	0.46	0.50	0.11
Interest Rate Change	−0.036	−0.148	−0.182**	−0.187***	−0.068	0.024
t-stat	0.35	1.47	2.14	2.45	1.24	0.30
Net Migration Rate Change	0.022***	0.025***	0.028***	0.030***	0.027***	0.030**
t-stat	3.15	3.54	3.08	3.33	3.33	2.52
Persons per House	0.009***	0.016***	0.018***	0.027***	0.015***	0.010***
t-stat	4.33	3.12	5.94	5.36	5.14	3.35
Share of Young	0.118***	0.106*	0.113*	0.072	0.123*	0.135**
t-stat	2.68	1.71	1.69	0.75	1.73	2.01
Housing Stock/GDP	−0.006*	−0.007**	−0.008**	−0.007**	−0.008***	−0.008***
t-stat	1.86	2.20	2.54	2.42	2.66	2.59
GDP per Capita Change	0.184	0.320	0.163	0.091	0.261	0.092
t-stat	0.71	1.54	0.95	0.85	1.46	0.79
CPI Inflation Exp.	0.000	−0.001	−0.001	−0.001	−0.001	0.000
t-stat	0.38	0.58	0.72	1.01	0.62	0.22
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,706	2,706	2,706	2,705	2,704	2,689
R2	0.147	0.140	0.134	0.140	0.135	0.133
F Statistic	15.58***	19.14***	13.26***	42.47***	25.48***	12.90***
RMSE	0.0517	0.0515	0.0517	0.0506	0.0490	0.0488

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

an increase in household size—which is strongly correlated with population density—is associated with higher residential investment.

Third, a larger existing housing stock acts as a break on new residential investment. As a corollary, countries with smaller housing stocks have seen faster growth in residential investment, an observation that is also consistent with the higher investment ratios observed in the earlier part of the sample (figure 1).

Fourth, the coefficients on GDP per capita are consistently positive, suggesting that higher income levels are associated with higher residential investment. From a statistical viewpoint, however, this effect is not significant. Nor are the mostly negative effects of PPI inflation (a proxy for construction costs) and expected consumer price inflation (a measure of macroeconomic uncertainty).[12]

The appendix shows the results for two variations of this baseline specification. When time fixed effects are omitted, the quantitative effects of interest rates and expected house price increases are stronger (table A.1). This is not entirely unexpected, given that time fixed effects are meant to capture global trends, such as the long-term decrease in interest rates and financial globalization.

When the time sample is limited to the post-2000 period, the responsiveness of residential investment to price variables increases (table A.2). This is not surprising, either. In the 1970s and 1980s, several countries had large government programs to increase housing supply. These naturally tended to be relatively unresponsive to market conditions. More importantly, many countries have seen financial liberalization since the 1980s, which resulted in increased availability of mortgages to households. Together, the lower share of housing financed by the government and the development of mortgage markets translated into greater sensitivity of residential investment to changes in house prices and financial conditions.[13]

[12]If run on a country-by-country basis, the explanatory power of the baseline model for six-quarter cumulative investment growth is largest for Spain ($R^2 = 0.73$), followed by France (0.58) and the United States (0.44). It is lowest for Sweden (0.20) and Australia (0.15), as shown in figure A.4.

[13]The more recent period was also characterized by heightened policy uncertainty in many economies, in particular after the GFC. Greater uncertainty could result in the postponement of construction projects due to uncertain returns. To address this issue, we used a measure from Baker, Bloom, and Davis (2016) that is based on the frequency of newspaper articles that discuss policy

4.2 Examining Asymmetries

Next we identify cyclical phases of residential investment in order to examine whether the effect of interest rates on residential investment is asymmetric across the cycle. We define residential investment upswings as periods when the quarterly change in the residential investment-to-GDP ratio is above the 75th percentile of the distribution (for each country) for at least one year. Thus, upswings are associated with a rapid increase in the share of residential investment in the economy. Similarly, residential investment downswings are defined as periods when the quarterly change in the residential investment-to-GDP ratio is below the 25th percentile of the distribution in each country. During such periods, the share of residential investment in the overall economy falls.

The cycles are computed using the four-quarter moving averages of the residential investment-to-GDP ratio to minimize the effect of temporary volatility in residential investment. Short upswings/downswings, i.e., those lasting less than four quarters, are not considered in order to have sufficient persistence in cyclical phases. Moreover, any gaps of less than four quarters between two identical phases (downswing or upswing) were eliminated, thus combining the two phases into a single one. The resulting 76 cyclical upswings and 65 downswings are shown in figures 5 and 6, while figure 7 shows the number of economies that are experiencing upswings and downswings at each point in time.

The earlier part of the sample saw a greater number of upswings in residential investment; since the mid-2000s, their occurrence has declined somewhat, only to rise again in the most recent period (figures 5 and 7). The longest upswing in the sample was the one in Spain, spanning the late 1990s and the early 2000s, when the share of residential investment in GDP rose by slightly more than 3 percentage points. Another recent case is Sweden, which went through an upswing in 2013–16, with the GDP share rising by 1.6 percentage points.

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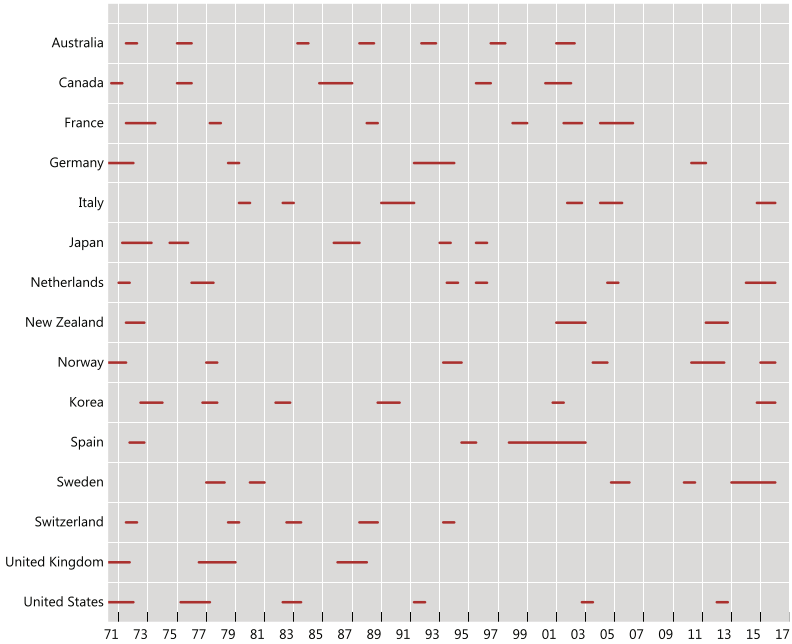
uncertainty. While we found that policy uncertainty is negatively associated with residential investment activity, the magnitude of the estimated effect was small. These results are available upon request.

Figure 5. Timeline of Upswings in Residential Investment, 1970–2016[a]



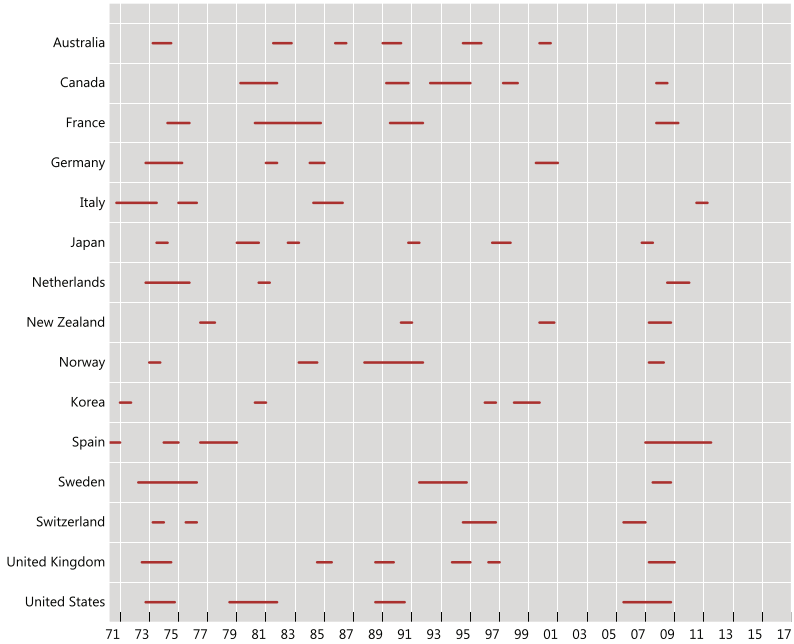
Sources: OECD, Economic Outlook database; national data; authors' calculations.

[a]Defined as quarterly growth (in percentage points) above the 75th percentile within each country, based on four-quarter moving averages of residential investment as a share of GDP. Short upswings lasting less than four quarters were dropped, and short gaps (less than four quarters between two upswings) were connected. For Switzerland, construction is used for residential investment.

Many countries saw downswings around the Great Financial Crisis (figures 6 and 7). Spain experienced a long investment slump during 2007–14, when the residential investment-to-GDP ratio fell by 5.5 percentage points. In the United States, the ratio declined by 3 percentage points in the aftermath of the GFC. In Norway, it fell by a similar amount in the late 1980s, and again in the early 1990s.

Next, we use the identified cyclical phases as zero-one dummy variables interacted with interest rates. We find that a change in interest rates by 100 basis points is associated with a decline in residential investment by 0.78 percent after four quarters during

Figure 6. Timeline of Downswings in Residential Investment, 1970–2016[a]

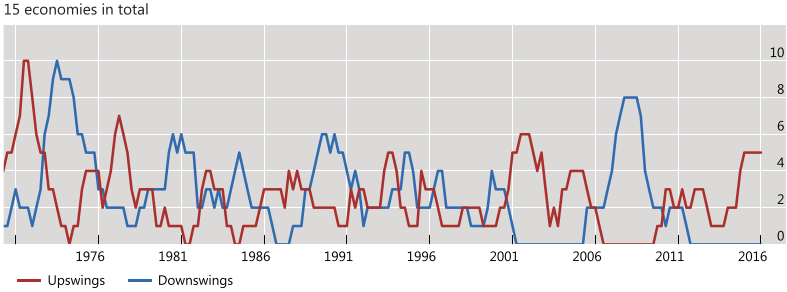


Sources: OECD, Economic Outlook database; national data; authors' calculations.
[a]Defined as quarterly growth (in percentage points) below the 25th percentile within each country, based on four-quarter moving averages of residential investment as share of GDP. Short downswings lasting less than four quarters were dropped, and short gaps (less than four quarters) between two downswings were connected. For Switzerland, construction is used for residential investment.

residential investment upswings (table 2), and with a bit less than half that during normal times (0.33 percent).[14] When we interact interest rate changes with downswings, we obtain much lower coefficient estimates (appendix table A.3). Intuitively, this asymmetry is not surprising: for a property developer, the real cost of borrowing is the interest rate minus the expected appreciation of the property's price. Thus, when house prices are expected to fall, the real cost

[14]These results use coefficient estimates that are statistically significant at a 10 percent level at least.

Figure 7. Number of Economies in Upswings and Downswings in Residential Investment



Sources: OECD, Economic Outlook database; national data; authors' calculations.
Notes: Upswings (downswings) are defined as quarterly growth, in percentage points, above the 75th percentile (below the 25th percentile) within each country, based on four-quarter moving averages of residential investment as share of GDP. Short swings lasting less than four quarters were dropped, and short gaps (less than four quarters) between two upswings were connected. For Switzerland, construction is used for residential investment.

of building may be too high even if interest rates were to drop to zero.[15]

We get similar results when we use quantile regressions instead of the above classification of upswings and downswings. The effects of interest rates on building activity are strongest when residential investment growth is peaking, and not statistically significant during a slump, as shown in table A.4 in the appendix.[16] This evidence is broadly in line with findings for aggregate output by Tenreyro and Thwaites (2016), among others.

One explanation for such dynamics is that borrowers become more sensitive to higher interest rates when residential investment is expanding rapidly. Both property developers and buyers may be incurring higher debt levels, and if the upturn coincides with a financial boom, marginal borrowers may benefit from greater availability of credit. We find some support for this narrative, as real total

[15]We thank an anonymous referee for this point.
[16]We thank an anonymous referee for suggesting this alternative test.

Table 2. Effects of Interest Rates during Upswings

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.054	−0.040	−0.025	0.003	0.018	0.010
t-stat	1.28	1.26	0.95	0.09	0.64	0.44
Real House Price Growth Exp.	0.449***	0.224**	0.093	0.053	−0.030	0.035
t-stat	4.93	2.36	1.09	0.50	0.35	0.44
Upswing	0.017***	0.010***	0.004**	−0.001	−0.003*	−0.005**
t-stat	5.71	5.07	2.05	0.45	1.67	2.37
Interest Rate Change	−0.011	−0.147	−0.152*	−0.173**	−0.045	−0.015
t-stat	0.10	1.36	1.92	2.24	0.72	0.19
Interest Rate Change*Upswing	−0.331	0.026	−0.451**	−0.224	−0.372	0.580
t-stat	1.44	0.12	2.33	0.88	1.52	1.02
Net Migration Rate Change	0.020***	0.024***	0.027***	0.030***	0.027***	0.031**
t-stat	3.35	3.94	3.41	3.34	3.37	2.57
Persons per House	0.010***	0.016***	0.018***	0.027***	0.015***	0.010***
t-stat	4.98	3.26	6.06	6.69	5.08	3.22
Share of Young	0.083***	0.086*	0.104*	0.075	0.129*	0.146**
t-stat	3.32	1.86	1.77	0.79	1.74	1.99
Housing Stock/GDP	−0.006***	−0.007**	−0.008***	−0.007**	−0.008***	−0.008**
t-stat	3.11	2.33	2.58	2.40	2.60	2.54
GDP per Capita Change	0.149	0.300	0.153	0.093	0.266	0.104
t-stat	0.58	1.49	0.93	0.85	1.45	0.91
CPI Inflation Exp.	0.000	0.000	−0.001	−0.001	−0.001	0.000
t-stat	0.21	0.49	0.67	1.01	0.64	0.29
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,706	2,706	2,706	2,705	2,704	2,689
R2	0.159	0.144	0.135	0.140	0.136	0.136
F Statistic	1,089.47***	54.22***	23.65***	311.00***	24.76***	24.14***
RMSE	0.0513	0.0514	0.0516	0.0506	0.0490	0.0487

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

credit growth is more than twice higher during residential investment upturns than downturns (median growth rates of 1.5 percent and 0.6 percent, respectively).

Does the relationship between interest rates and residential investment differ for interest rate increases and decreases? Previous research has highlighted the asymmetric effects of monetary contractions and expansions on economic activity, with interest rate hikes generally having stronger effects than interest rate cuts (e.g., Angrist, Jordà, and Kuersteiner 2013). Table 3 concurs with that evidence, as interest rate increases are found to be associated with lower residential investment, over multiple horizons, with economically and statistically significant coefficients. In contrast, and perhaps surprisingly, the correlation of residential investment with interest rate declines is not statistically significant at any horizon. This result is not driven by the Great Financial Crisis and its aftermath, as additional estimates show that similar dynamics prevailed from 1970 to 2006.[17]

Why would interest rate increases have greater effects on residential investment than interest rate decreases? A key reason could be downward rigidity in real house prices. Sellers may be unwilling to let prices fall sufficiently to generate expectations of new price increases. Figure 8 supports this conjecture, showing that house prices rose much more strongly (1.5 percent quarter on quarter) during median residential investment upswings than they fell during median downswings (−0.6 percent). Another reason is that the tightening and easing of financial conditions have asymmetric effects on creditworthiness of property builders. The default rate of unhedged property developers tends to rise after an interest rate increase, but is bounded when interest rates decrease, as the default rate cannot fall below zero. A third reason is that the pass-through of changes in interest rates from financial intermediaries to final users of funds might be larger when interest rates rise than when they fall.

Construction value added relative to GDP also displays asymmetry between upswings and downswings: during a median upswing, it rises by less than 0.1 percentage point a year on average, while during a median downswing it falls by close to 0.3 percentage point.

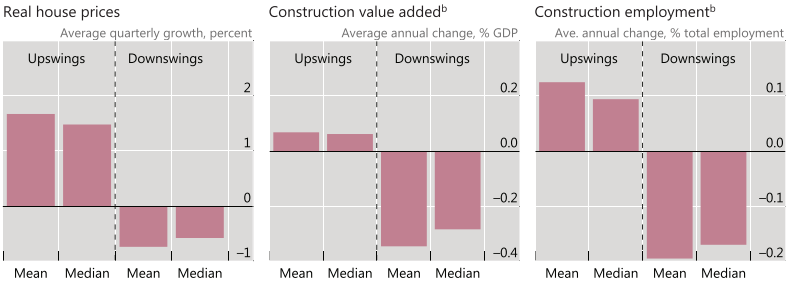
[17]These results are available upon request.

Table 3. Testing for Asymmetries: Rate Hikes versus Rate Reductions

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.049	−0.037	−0.024	0.001	0.016	0.009
t-stat	1.09	1.10	0.90	0.05	0.57	0.43
Real House Price Growth Exp.	0.530***	0.271***	0.113	0.040	−0.050	0.005
t-stat	5.30	2.88	1.29	0.42	0.55	0.07
Interest Rate Increase	−0.309**	−0.335	−0.294***	−0.418***	−0.228**	−0.115
t-stat	2.53	1.45	4.15	3.77	2.00	0.90
Interest Rate Decrease	0.262	0.056	−0.058	0.067	0.107	0.175
t-stat	1.23	0.43	0.40	0.54	0.80	1.38
Net Migration Rate Change	0.022***	0.024***	0.028***	0.030***	0.026***	0.030***
t-stat	3.08	3.49	3.06	3.29	3.30	2.50
Persons per House	0.009***	0.016***	0.018***	0.027***	0.016***	0.010***
t-stat	4.56	3.98	6.00	6.80	5.23	3.43
Share of Young	0.107***	0.099*	0.109*	0.063	0.117*	0.130**
t-stat	2.84	1.91	1.79	0.73	1.79	2.09
Housing Stock/GDP	−0.006**	−0.007**	−0.008***	−0.008***	−0.008***	−0.008***
t-stat	2.05	2.33	2.62	2.58	2.77	2.69
GDP per Capita Change	0.151	0.298	0.149	0.063	0.241	0.075
t-stat	0.57	1.45	0.88	0.56	1.39	0.65
CPI Inflation Exp.	0.000	0.000	−0.001	−0.001	0.000	0.000
t-stat	0.16	0.44	0.63	0.82	0.49	0.11
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,706	2,706	2,706	2,705	2,704	2,689
R2	0.150	0.141	0.134	0.142	0.136	0.134
F Statistic	29.65***	33.99***	15.95***	56.50***	29.65***	11.10***
RMSE	0.0516	0.0515	0.0517	0.0506	0.0490	0.0487

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Figure 8. House Prices, Construction Output, and Employment during Residential Investment Upswings and Downswings[a]



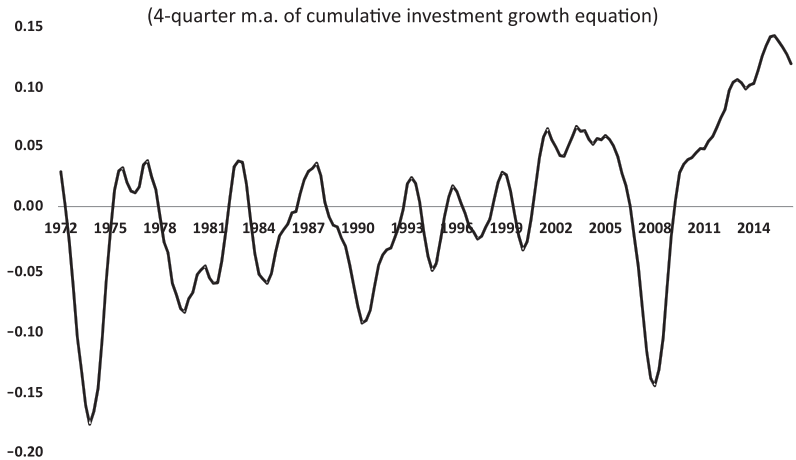
Sources: OECD; Datastream; national data; authors' calculations.
[a]See figures 5 and 6 for definitions of upswings and downswings.
[b]When two or more quarters within a calendar year are upswing or downswing quarters, the whole year is considered an upswing or downswing period.

Similarly, the share of employment in the construction sector rises less during a median upswing (0.09 percentage point a year on average) than it falls during a median downswing (−0.17 percentage point).

These findings corroborate previous evidence that the bulk of adjustment during housing downswings occurs through volumes rather than prices (Poterba 1984; Leamer 2015). In contrast, during upswings, labor shortages and supply and administrative bottlenecks make it more difficult to expand construction, so house prices rise more than housing volumes (Corder 2008).[18]

[18]In additional estimations focusing only on upswings, we find that interest rate reductions increase residential investment growth by less in absolute magnitude than interest rate increases reduce it, consistent with adjustment costs that make it difficult to expand construction. We also test the robustness of the findings in figure 8 to two other definitions of upswings and downswings. First, we construct the 25th and 75th percentiles based on the entire sample distribution (rather than the country-specific distributions) of the change in residential investment-to-GDP ratios. Second, we use the growth in real residential investment, rather than the change in its GDP share, to compute the 25th and 75th percentiles. The asymmetries identified in figure 8 are robust to these alternative definitions. These results are available upon request.

Figure 9. Estimated Global Factor in Residential Investment Growth



Upswings and downswings in residential investment are partly synchronized across countries, suggesting that there may be some important common global factors in their dynamics. The quarterly time fixed effects can be regarded as a proxy for "global" residential investment growth, after controlling for country fixed effects and other determinants of residential investment. Figure 9 shows that there are three periods during which time fixed effects turned strongly negative. The first was in the early 1970s, in the immediate aftermath of the first oil shock, when the Federal Reserve tightened monetary conditions aggressively. The second was in the early 1990s, when global credit contracted sharply. Finally, the time dummies turned strongly negative between the second half of 2007 and early 2010, during the GFC.

The inclusion of time fixed effects stacks the cards against finding significant effects from other explanatory variables. Indeed, if time fixed effects are omitted, the coefficients on other explanatory variables increase in statistical significance (appendix table A.1). We nevertheless include time fixed effects in our baseline estimations given the higher explanatory power of the models that include them and the possibility that they capture relevant and otherwise omitted global conditions.

Finally, residential investment dynamics could be affected by credit market developments.[19] Indeed, when we divide the sample according to whether the International Monetary Fund's mortgage market development index (MMI) is above or below the sample median, we find that the response of residential investment to interest rates and expected house prices is larger in countries with more developed mortgage markets (last column of table 4). We also find indicative evidence on a smaller sample size that higher credit spreads inhibit residential investment (table A.5 in the appendix).

5. Evidence from Recession-Prediction Models

The finding in section 2 that downturns in the housing cycle precede those in GDP can be formally tested for its recession-prediction properties. We build on Estrella and Hardouvelis (1991) and Rudebusch and Williams (2009), who presented evidence that the yield curve was the best single predictor of recessions. Using logistic regressions for the nine economies in our sample for which long time series on the yield curve are available (at least from 1994 onward), we found that information on residential investment growth consistently improved the performance of the yield curve as a predictor of recessions across countries.

Column 1 of table 5 confirms the result in previous literature that the term spread provides useful information for forecasting the start of a recession (defined as two quarters of consecutive negative growth) during the following four quarters. Column 2 shows that the incidence of negative residential investment growth over the past four quarters also helps predict a recession. Notably, the statistical significance of both variables remains high when they are included in the same regression, as shown in column 3. In such a model, the pseudo R^2 more than doubles relative to the specification that features the yield curve slope alone (column 1). Strikingly, column 5 shows that business—unlike residential—investment declines are not very useful on their own for predicting recessions, and only add marginal prediction value to the benchmark model with the yield curve slope. Current GDP growth cannot match the forecasting properties

[19]Global credit market developments or secular trends are inevitably captured in the time dummies.

Table 4. Results by Mortgage Market Development

Dependent Variable: Residential Investment Growth (Real, Log Differences)		
	Cumulative Growth between Quarter t and $t+6$	
	MMI < 0.575	**MMI > 0.575**
PPI Inflation	−0.523***	0.034
t-stat	3.90	0.33
Real House Price Growth Exp.	0.413*	1.014***
t-stat	1.81	3.52
Interest Rate Change	−0.412**	−0.753**
t-stat	2.24	2.48
Net Migration Rate Change	0.184***	0.178***
t-stat	8.78	5.94
Persons per House	0.079**	0.089***
t-stat	2.39	4.25
Share of Young	0.610***	0.042
t-stat	3.35	0.11
Housing Stock/GDP	−0.023***	−0.060***
t-stat	2.90	4.02
GDP per Capita Change	0.869**	1.415***
t-stat	2.13	3.16
CPI Inflation Exp.	0.003	−0.006***
t-stat	1.59	2.92
Country Fixed Effects	Yes	Yes
Time Fixed Effects	Yes	Yes
Observations	1,082	1,607
R2	0.441	0.317
F Statistic	20.27***	15.28***
RMSE	0.0775	0.1465

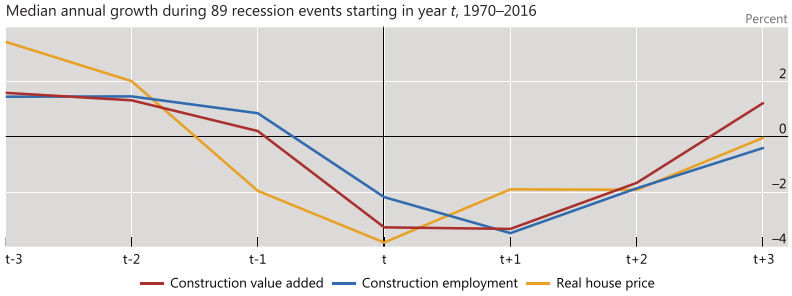
Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Table 5. Logistic Regressions for Probability of Recession Starting within Next Four Quarters

Probability of Start of Recession between $t+1$ and $t+4$ (Log Odds-Ratio)							
	Residential Investment				Business Inv.	GDP Growth	HP Growth
	(1)	(2)	(3)	(4)	(5)	(6)	(7)
Yield Curve Slope (10 y − 3 m)	-0.330**		−0.264***	−0.259***	−0.342***	−0.329***	−0.277***
t-stat	5.08		4.01	3.93	5.26	5.07	3.96
Number of Quarters of Negative *Residential* Investment Growth between $t-3$ and t		0.581***	0.548***				
t-stat		7.35	6.76				
Number of Quarters of Negative *Business* Investment Growth between $t-3$ and t					0.146*		
t-stat					1.82		
Negative Residential Investment Growth in t				0.683***			
t-stat				3.45			
Negative Residential Investment Growth in $t-1$				0.488**			
t-stat				2.42			
Negative Residential Investment Growth in $t-2$				0.660***			
t-stat				3.32			
Negative Residential Investment Growth in $t-3$				0.354*			
t-stat				1.81			
GDP Growth, Annual ($t-4$ to t)						0.024	
t-stat						0.19	
Resid. House Price Growth, Annual ($t-4$ to t)							−0.199***
t-stat							3.75
Observations	941	941	941	941	941	941	926
Number of Countries	9	9	9	9	9	9	9
Pseudo R2	0.034	0.077	0.098	0.101	0.038	0.034	0.061
Chi-Square	25.6	57.5	73.8	75.4	28.8	25.6	43.5
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes	Yes

Notes: ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, CH, DE, ES, IT, JP, GB, and US. Inclusion based on availability of all time series above since at least 1994. Recession quarters after starting period dropped from the sample.

Figure 10. Construction-Sector and Real House Price Growth during Recessions



Sources: OECD; national data; BIS Residential Property Price Statistics; authors' calculations.
Note: Recession events are defined as negative annual GDP growth in real terms.

of residential investment, either (column 6). House price variation, however, does add value, even though its significance is below that for residential investment growth (column 7).

To further illustrate these findings, we plot in appendix figure A.5 the recession probabilities obtained from the parsimonious specification 3 for Australia, the United Kingdom, and the United States. The strong rise in estimated recession probabilities prior to downturns is particularly noticeable for the United States but can also be seen for the recessions Australia and the United Kingdom experienced in the early 1990s.

Through which channels could residential investment contribute to output recessions? Figure 10 plots median growth in construction value added, construction-sector employment, and real house prices over the business cycle. As the data related to the construction sector are annual, recessions in this graph are defined simply as years of negative real GDP growth (periods on the horizontal axis denote years).[20] The first turn is noted in house prices: real house price

[20]Figure 10 uses annual data due to a lack of higher-frequency data on the construction sector. When we use quarterly data on residential investment for the economies included in the recession prediction exercise, we find that growth in residential investment becomes negative five quarters before the start of a recession.

growth decelerates two years before a recession and then becomes negative (–2 percent) in the year preceding the output decline. It then dips further and remains negative overall for four consecutive years. Construction value added then decelerates and drops by 3 percent in the year the recession starts. Employment in the sector follows the same path and falls by around 2 percent in the year when GDP growth turns negative. In addition to these real activity channels, wealth and credit collateral channels induced by the fall in house prices may also restrain activity (e.g., Campbell and Cocco 2007). House price declines may also sap consumer confidence, restraining private consumption.

6. Robustness Checks

We performed a number of additional robustness checks. First we replaced residential investment growth with its GDP share as the dependent variable. The results (not shown) did not change the above findings.

Next we excluded the years of the Great Financial Crisis (2007–09) from the sample. During those three years, all countries in the sample saw a decline in the residential investment-to-GDP ratio, ranging from 0.3 percentage in Italy and Switzerland to 3 percentage points in Spain. Appendix table A.6 shows, however, that the results largely remain robust to the exclusion of the GFC years.

We have also examined whether the results are driven by individual economies with outlier observations. In these tests, we found that the results are robust to excluding either Korea or Italy, i.e., the economies with the highest and lowest average residential investment growth rates, respectively, in our sample.[21]

Finally, we evaluated the relevance of public housing for the results. Our estimated model, including determinants such as construction costs and income levels, is more relevant for economies where public-sector residential investment is unimportant. One should recognize, however, that variables such as real house price growth are relevant for both private and public housing provision,

[21] These results are available upon request.

through a higher Tobin's q and the objective of guaranteeing adequate supply of low-cost housing for lower-income groups. To capture the impact of public housing, we estimated the baseline model excluding all countries where social rental dwellings accounted for at least 10 percent of the total housing stock in 2000 and/or 2015 (France, the Netherlands, and the United Kingdom, based on data from the OECD's Affordable Housing Database). As one would expect, the fit of the model improves a bit when countries with the highest public housing provision are excluded (appendix table A.7). More importantly, the sensitivity of residential investment to real house prices and interest rates increases.

7. Conclusions

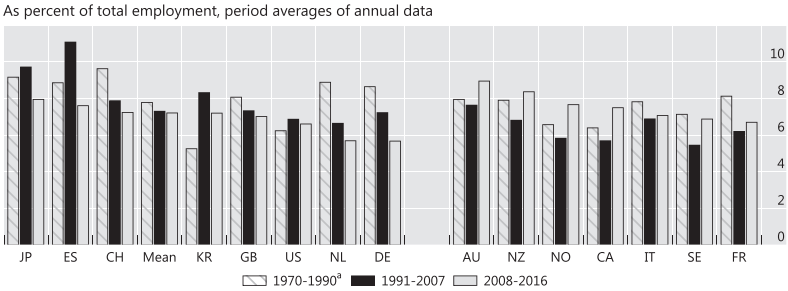
In this paper we analyzed the behavior and main drivers of residential investment using a panel data set for 15 advanced economies since the beginning of the 1970s. Our estimations suggest that house price growth, net migration inflows, and the size of the existing housing stock are the most important drivers of residential investment. We further show that interest rate increases affect residential investment more than interest rate decreases, and that interest rate changes have larger effects on residential investment when its share in GDP is rising strongly.

Also, residential investment consistently anticipates economic downturns across countries. Adding information on residential investment dynamics noticeably improves the performance of standard recession-prediction models.

One interesting issue for future research is the relevance of country-specific institutional factors such as housing supply regulations for residential investment. Such factors cannot be incorporated easily in a cross-country study but they could matter for country-specific investment dynamics. Another interesting issue relates to the use of macroprudential policies directed at the housing sector, which have gained prominence in advanced economies over the past decade. To the extent that they become widely used, macroprudential policies may also affect the short-term dynamics of residential investment activity in the future.

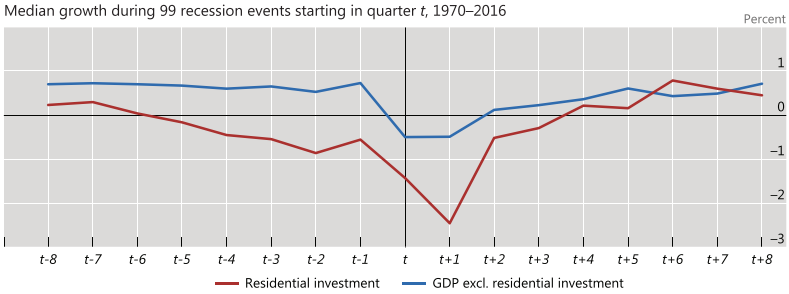
Appendix

Figure A.1. Construction Employment Generally Highest before the 1990s, Fell Post-Crisis



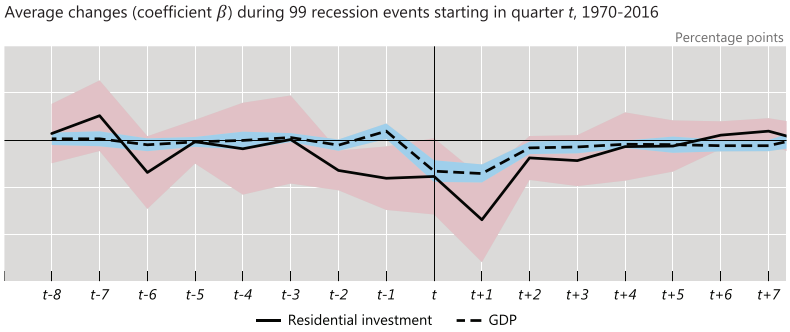
Sources: OECD; Datastream; national data; authors' calculations.
[a]For New Zealand, data are available from 1978.

Figure A.2. Residential Investment, and GDP Excluding Residential Investment, during Recessions



Sources: OECD, Economic Outlook database; national data; authors' calculations.
Note: Recession events are defined as at least two consecutive quarters of negative GDP growth based on seasonally adjusted data.

Figure A.3. Residential Investment versus GDP Growth during Recessions



Sources: OECD, Economic Outlook database; national data; authors' calculations.
Notes: This graph estimates the average changes of residential investment and GDP growth during eight quarters before and after 99 recession events relative to normal times, controlling for country and time fixed effects. The regression is of the following type:

$$IHV \text{ or } GDP \text{ } growth_{i=country,\ t=time,\ rec=start\ of\ recession}$$
$$= \sum_{k=-8}^{8} \beta_k 1_{t=rec+k} + \gamma_i + \delta_t + \varepsilon_{i,t,rec}.$$

The ranges indicate the 95 percent confidence intervals.

Figure A.4. Explanatory Power of Baseline Model by Countries: R^2s

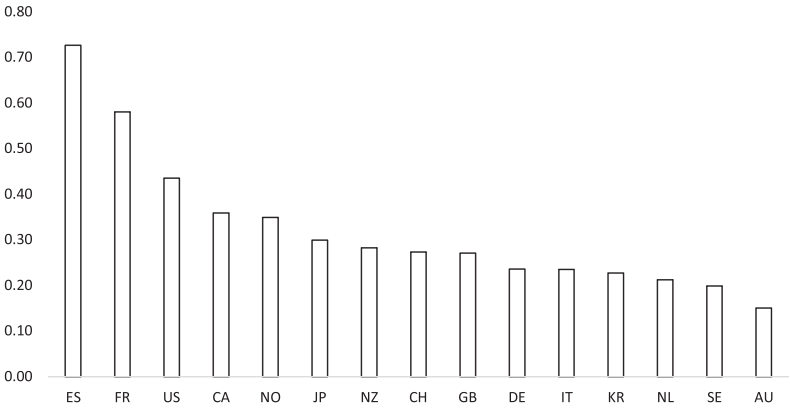


Figure A.5. Estimated Recession Probabilities

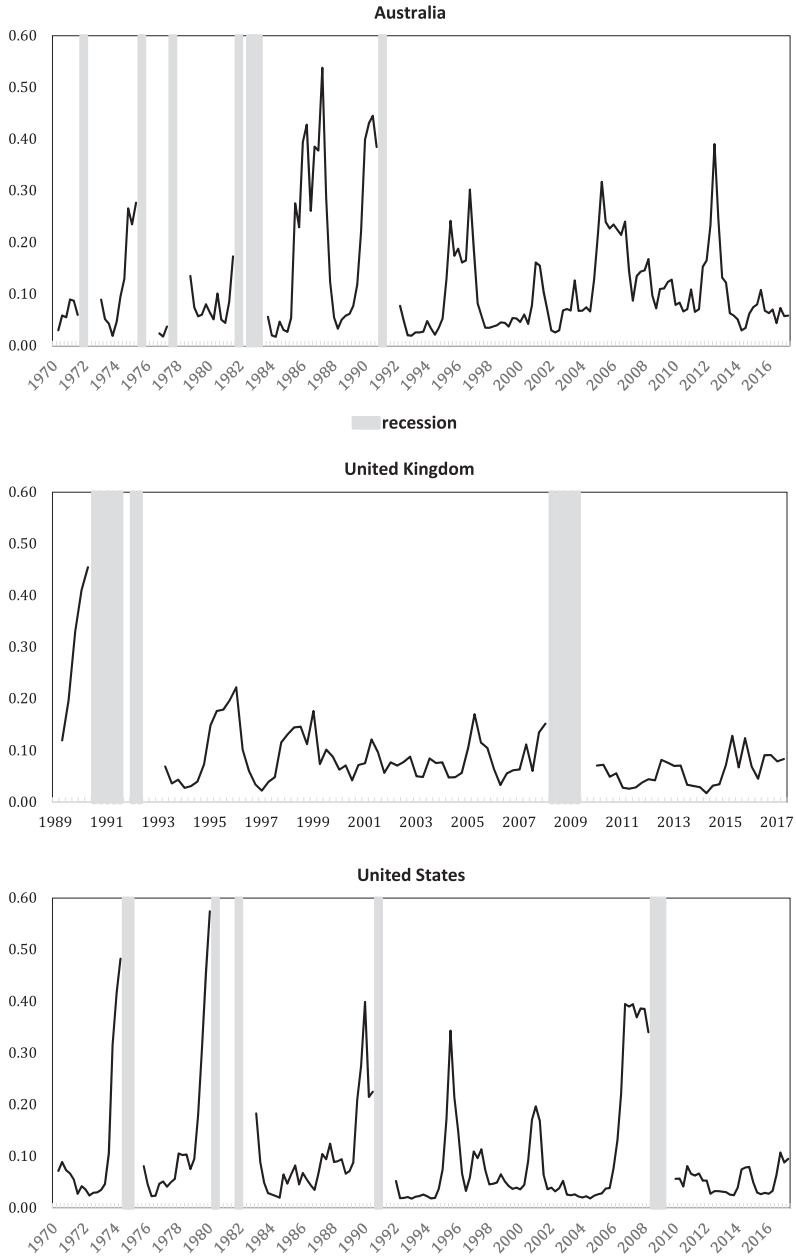


Table A.1. Specification without Time Fixed Effects

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.103**	−0.095***	−0.078***	−0.050**	−0.035	−0.033
t-stat	2.25	2.70	2.79	2.51	1.17	1.31
Real House Price Growth Exp.	0.584***	0.314***	0.101	0.020	−0.080	−0.021
t-stat	6.56	3.82	1.18	0.19	0.80	0.23
Interest Rate Change	−0.140	−0.275**	−0.248**	−0.211**	−0.065	−0.069
t-stat	1.09	2.31	2.07	1.99	1.13	1.06
Net Migration Rate Change	0.021***	0.025***	0.029***	0.030***	0.029***	0.029***
t-stat	3.06	4.12	3.60	3.80	3.57	2.62
Persons per House	0.009***	0.013*	0.014**	0.018**	0.005	−0.001
t-stat	2.99	1.80	2.31	2.28	0.88	0.19
Share of Young	−0.012	−0.047	−0.049	−0.084	−0.036	−0.014
t-stat	0.37	1.17	1.09	1.15	0.73	0.33
Housing Stock/GDP	−0.005*	−0.006*	−0.007**	−0.006***	−0.006**	−0.007**
t-stat	1.74	1.84	2.21	2.96	2.15	2.28
GDP per Capita Change	0.261	0.220	−0.009	−0.054	0.012	−0.080
t-stat	1.09	1.11	0.06	0.45	0.08	0.74
CPI Inflation Exp.	0.000	0.000	−0.001	−0.001	0.000	0.000
t-stat	0.11	0.30	0.57	0.84	0.42	0.08
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,706	2,706	2,706	2,705	2,704	2,689
R2	0.058	0.042	0.029	0.028	0.017	0016

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Table A.2. Post-2000 Sample

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.005	0.015	0.047	0.043*	0.006	−0.036
t-stat	0.12	0.41	1.55	1.71	0.18	0.96
Real House Price Growth Exp.	0.700***	0.614***	0.462***	0.176	0.167	0.017
t-stat	4.73	4.12	3.75	1.42	1.19	0.10
Interest Rate Change	−0.488	−1.133*	−0.975**	−1.205**	−1.105***	−1.329*
t-stat	1.20	1.74	2.08	2.47	2.67	1.78
Net Migration Rate Change	0.011	0.011	0.014	0.012	0.003	0.002
t-stat	1.34	1.25	1.13	1.02	0.22	0.11
Persons per House	−0.040	−0.011	−0.021	−0.044	−0.044*	−0.042
t-stat	1.55	0.34	0.65	1.62	1.77	1.56
Share of Young	0.401***	0.274	0.358*	0.460**	0.392**	0.358*
t-stat	2.62	1.47	1.81	2.50	2.29	1.83
Housing Stock/GDP	−0.016***	−0.016***	−0.021***	−0.023***	−0.027***	−0.028***
t-stat	3.15	2.74	3.51	3.84	4.50	4.74
GDP per Capita Change	0.272	0.421	−0.222	−0.066	−0.032	−0.030
t-stat	1.04	0.98	0.68	0.42	0.16	0.10
CPI Inflation Exp.	−0.003**	−0.001	−0.002	−0.003	−0.001	0.000
t-stat	2.57	0.58	0.81	1.30	0.35	0.14
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	1,020	1,020	1,020	1,019	1,018	1,003
R2	0.250	0.253	0.249	0.252	0.255	0.259

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Table A.3. Downswings

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.045	−0.036	−0.023	0.003	0.017	0.009
t-stat	1.03	1.07	0.90	0.12	0.61	0.42
Real House Price Growth Exp.	0.433***	0.212**	0.103	0.077	−0.025	0.021
t-stat	4.38	2.26	1.22	0.74	0.32	0.27
Downswing	−0.017***	−0.010**	−0.002	0.004	0.003	0.002
t-stat	4.13	2.43	0.53	0.86	0.76	0.48
Interest Rate Change	−0.259	−0.113	−0.253*	−0.432***	−0.202**	0.002
t-stat	1.17	0.78	1.80	3.57	2.33	0.01
Interest Rate Change*Downswing	0.303	−0.064	0.102	0.363**	0.200*	0.034
t-stat	1.31	0.39	0.56	2.54	1.90	0.21
Net Migration Rate Change	0.018***	0.023***	0.027***	0.030***	0.027***	0.031**
t-stat	2.58	3.28	3.01	3.34	3.35	2.55
Persons per House	0.010***	0.016***	0.018***	0.027***	0.015***	0.010***
t-stat	4.90	3.22	6.00	6.69	5.12	3.32
Share of Young	0.111***	0.104*	0.112*	0.070	0.121*	0.135*
t-stat	2.64	1.77	1.72	0.73	1.71	2.02
Housing Stock/GDP	−0.005***	−0.006**	−0.008**	−0.007**	−0.008***	−0.008***
t-stat	2.67	2.16	2.53	2.41	2.66	2.60
GDP per Capita Change	0.155	0.298	0.161	0.108	0.271	0.097
t-stat	0.61	1.43	0.93	0.99	1.45	0.81
CPI Inflation Exp.	0.000	−0.001	−0.001	−0.001	−0.001	0.000
t-stat	0.36	0.60	0.71	0.97	0.60	0.22
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,706	2,706	2,706	2,705	2,704	2,689
R2	0.159	0.144	0.134	0.142	0.136	0.134

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Table A.4. Quantile Regression

Dependent Variable: Residential Investment Growth (Real, Log Differences)			
	Cumulative Growth between Quarter t and $t+6$		
	5th Quantile	**50th Quantile**	**95th Quantile**
PPI Inflation	−0.236	−0.338***	−0.383**
t-stat	1.16	4.02	2.52
Real House Price Growth Exp.	1.844***	0.947***	0.479
t-stat	4.59	4.88	1.36
Interest Rate Change	−0.592	−0.895***	−1.291***
t-stat	1.50	2.79	3.45
Net Migration Rate Change	0.156***	0.152***	0.189***
t-stat	3.81	8.47	6.53
Persons per House	−0.051***	0.037*	0.142***
t-stat	3.21	1.88	4.30
Share of Young	0.322	−0.201	0.144
t-stat	1.28	1.34	0.56
Housing Stock/GDP	−0.093***	−0.029***	−0.001
t-stat	5.48	3.65	0.09
GDP per Capita Change	−0.317	0.394	−0.661
t-stat	0.56	1.38	1.18
CPI Inflation Exp.	−0.007**	−0.003***	0.000
t-stat	2.28	3.05	0.02
Country Fixed Effects	Yes	Yes	Yes
Observations	2,689	2,689	2,689
Pseudo R2	0.200	0.077	0.252

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on bootstrapping, with 100 replications. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Table A.5. Including Credit Spreads

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$	$t+2$	$t+3$	$t+4$	$t+5$	$t+6$
	(1)	(2)	(3)	(4)	(5)	(6)
PPI Inflation	0.013	0.024	0.043	0.036	−0.001	−0.039
t-stat	0.37	0.65	1.25	1.35	0.03	1.14
Real House Price Growth Exp.	0.823***	0.572***	0.536***	0.249*	0.301***	0.206*
t-stat	5.45	3.94	3.23	1.83	3.58	1.87
Interest Rate Change	−0.937**	−0.631	−0.736**	−0.742	−0.793*	−1.446*
t-stat	2.39	1.06	2.20	1.56	1.91	1.67
Corporate Spread Change	−1.062***	0.706	0.623	−0.126	0.415	−0.900***
t-stat	2.66	1.18	0.89	0.31	1.04	2.25
Net Migration Rate Change	−0.056	−0.069	−0.082	−0.114	−0.092	−0.091
t-stat	0.16	0.41	0.05	0.16	0.42	1.03
Persons per House	0.360**	0.479**	0.538***	0.713***	0.544***	0.555***
t-stat	2.24	2.30	2.56	4.06	2.97	2.95
Share of Young	−0.016**	−0.019**	−0.023***	−0.027***	−0.029***	−0.031***
t-stat	2.46	2.52	2.60	3.90	3.30	3.19
Housing Stock/GDP	0.384**	0.155***	0.068***	0.002***	−0.034***	0.201***
t-stat	2.22	2.74	3.32	3.79	4.18	5.15
GDP per Capita Change	−0.002	−0.003	−0.003	−0.004	−0.001	−0.001
t-stat	1.33	0.65	0.26	0.01	0.15	1.00
CPI Inflation Exp.	−0.011	0.007*	0.006*	−0.001**	0.004	−0.009
t-stat	0.92	1.73	1.70	2.20	0.58	0.31
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	882	882	882	881	880	866
R2	0.306	0.285	0.285	0.279	0.271	0.288

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NO, ES, SE, CH, GB, and US.

Table A.6. Sample without GFC

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$	$t+2$	$t+3$	$t+4$	$t+5$	$t+6$
	(1)	(2)	(3)	(4)	(5)	(6)
PPI Inflation	−0.059	−0.049	−0.034	−0.008	0.015	0.020
t-stat	1.13	1.39	1.22	0.26	0.51	1.01
Real House Price Growth Exp.	0.496***	0.244**	0.088	0.020	−0.042	0.020
t-stat	4.56	2.49	1.03	0.20	0.47	0.26
Interest Rate Change	−0.029	−0.146	−0.179**	−0.188**	−0.067	0.033
t-stat	0.28	1.44	2.08	2.41	1.20	0.44
Net Migration Rate Change	0.022***	0.025***	0.029***	0.031***	0.029***	0.032***
t-stat	2.76	3.55	3.19	3.49	4.09	2.66
Persons per House	0.009***	0.017***	0.018***	0.027***	0.014***	0.007*
t-stat	4.58	4.22	6.01	6.75	4.71	1.82
Share of Young	0.124***	0.110*	0.110	0.059	0.110	0.127*
t-stat	2.64	1.69	1.53	0.57	1.50	1.87
Housing Stock/GDP	−0.007**	−0.007**	−0.007**	−0.007	−0.006**	−0.006**
t-stat	2.25	2.43	2.48	1.64	2.02	2.07
GDP per Capita Change	0.177	0.304	0.168	0.107	0.282	0.083
t-stat	0.64	1.46	1.04	0.92	1.38	0.73
CPI Inflation Exp.	−0.000	−0.001	−0.001	−0.001	−0.001	0.000
t-stat	0.34	0.50	0.64	0.91	0.59	0.31
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,526	2,526	2,526	2,525	2,524	2,509
R2	0.130	0.124	0.119	0.125	0.120	0.120

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, FR, DE, IT, JP, KR, NL, NZ, NO, ES, SE, CH, GB, and US.

Table A.7. Without Countries Where Social Dwelling Is Above 10 Percent of Housing Stock

Dependent Variable: Residential Investment Growth (Real, Log Differences)						
	Growth in Quarter $t+h$					
	$t+1$ (1)	$t+2$ (2)	$t+3$ (3)	$t+4$ (4)	$t+5$ (5)	$t+6$ (6)
PPI Inflation	−0.062	−0.055	−0.030	−0.008	0.019	0.018
t-stat	1.13	1.31	0.95	0.26	0.70	1.01
Real House Price Growth Exp.	0.615***	0.284***	0.203**	0.151*	−0.033	−0.013
t-stat	5.64	2.84	2.07	1.68	0.27	0.13
Interest Rate Change	0.020	−0.213**	−0.164*	−0.169**	−0.108*	0.009
t-stat	0.25	2.20	1.90	2.06	1.86	0.13
Net Migration Rate Change	0.023***	0.028***	0.030***	0.033***	0.031***	0.033***
t-stat	3.22	4.05	3.33	3.62	5.20	2.96
Persons per House	0.008***	0.019***	0.019***	0.028***	0.017***	0.012***
t-stat	4.10	4.70	6.30	7.05	5.71	4.14
Share of Young	0.145***	0.136**	0.147**	0.104	0.154**	0.165**
t-stat	3.72	2.22	2.29	1.06	1.97	2.14
Housing Stock/GDP	−0.005*	−0.006	−0.008*	−0.008*	−0.008*	−0.008*
t-stat	1.82	1.64	1.88	1.89	1.90	1.93
GDP per Capita Change	0.395***	0.152	0.182	0.057	0.262	0.080
t-stat	3.47	0.69	1.03	0.59	1.26	0.67
CPI Inflation Exp.	−0.000	−0.001	−0.001	−0.001	−0.001	−0.001
t-stat	0.27	0.68	0.77	1.04	0.80	0.66
Country Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Time Fixed Effects	Yes	Yes	Yes	Yes	Yes	Yes
Observations	2,160	2,160	2,160	2,159	2,158	2,146
R2	0.169	0.152	0.147	0.153	0.148	0.145

Notes: Time span is from 1970:Q1 to 2017:Q2. All right-hand-side variables are at period t. Reported t-statistics below coefficients are based on robust standard errors clustered by country. ***, **, and * denote statistical significance at 1 percent, 5 percent, and 10 percent, respectively. Countries covered are AU, CA, DE, IT, JP, KR, NZ, NO, ES, SE, CH, and US.

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Monetary Policy as an Optimum Currency Area Criterion*

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The costs and benefits of moving from a flexible exchange rate regime to a monetary union depend critically on the conduct of monetary policy. In particular, whether countries are better off in one or the other currency regime is sensitive not only to the choice of the variables that monetary policy targets but also to the strength of the response to these target variables. In addition to being an optimum currency area (OCA) criterion itself, monetary policy can modify the nature of traditional OCA criteria, such as the degree of trade openness.

JEL Codes: F33, F41, E52.

1. Introduction

Over the decades since its initiation by Mundell (1961), the optimum currency area (OCA) theory has identified numerous criteria that are considered important in determining whether countries benefit from monetary unification. Traditional OCA criteria include the degree of labor mobility, price and wage flexibility, trade openness, the incidence of asymmetric shocks, country size, the similarity of economic structures, the degree of product diversification, and the degree of fiscal integration.[1]

However, one criterion has received hardly any attention, although it is critical for the welfare implications of monetary unification: the conduct of monetary policy. In particular, I show that whether countries are better off under a flexible exchange rate regime or a monetary union is sensitive not only to the choice of the variables that monetary policy targets but also to the strength of the response

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[1]Excellent surveys of the OCA literature are Mongelli (2002), Dellas and Tavlas (2009), Beetsma and Giuliodori (2010), and De Grauwe (2012).

to these target variables. When monetary policy in each country responds to inflation aggressively or implements a high degree of interest rate smoothing, forming a monetary union, where the common monetary authority continues to follow the same policy, tends to make countries worse off in terms of welfare by reducing macroeconomic stability. By contrast, when monetary policy responds to inflation only modestly or implements a low degree of interest rate smoothing, forming a monetary union with exactly the same monetary policy tends to make countries better off. Furthermore, monetary unification is beneficial when monetary policy responds to *output*, whereas it is costly when monetary policy responds to the *output gap*. And finally, it is important whether countries respond to the nominal exchange rate and whether they do so in a coordinated or uncoordinated way. Monetary unification is generally beneficial in the latter case, but not in the former case.

I show that monetary policy, in addition to being an OCA criterion itself, has the potential to modify the nature of traditional OCA criteria, such as the degree of trade openness. Whether the likelihood of a monetary union being beneficial *increases* with the degree of trade openness, as proposed by the vast bulk of OCA studies, depends critically on whether independent monetary policy targets producer price inflation or consumer price inflation. In the former case, it is also possible that the likelihood of a monetary union being beneficial *decreases* with the degree of trade openness.

The conduct of monetary policy matters for the welfare implications of monetary unification for two reasons. First, monetary policy determines to what extent a flexible nominal exchange rate fosters or hampers macroeconomic stabilization, even if monetary policy does not target the nominal exchange rate explicitly. A flexible nominal exchange rate renders monetary policy more powerful in the sense that monetary policy affects all welfare-relevant variables directly. By contrast, in a monetary union, the influence of monetary policy is limited by the fixed exchange rate, especially with respect to international relative prices such as the terms of trade. However, the fact that monetary policy is more powerful under a flexible exchange rate regime is a double-edged sword. When the interest rate policy happens to move the nominal exchange rate in the "right" direction, forming a monetary union generally—not always (see second reason)—reduces macroeconomic stability and welfare by

eliminating the *stabilizing* effects of the nominal exchange rate. By contrast, when the interest rate policy happens to move the nominal exchange rate in the "wrong" direction, forming a monetary union increases macroeconomic stability and welfare by eliminating the *destabilizing* effects of the nominal exchange rate. Importantly, which policies move the exchange rate in which direction is anything but obvious.

The second reason for monetary policy being an OCA criterion is the existence of a benefit that is inherent to monetary unions (Groll and Monacelli 2020). This renders a monetary union beneficial even for interest rate policies that move the nominal exchange rate in the right direction, e.g., a modest response to inflation. While constraining monetary policy to some extent, the fixed exchange rate has the advantage of stabilizing private-sector expectations about future inflation and thereby stabilizing actual inflation. This can overcompensate for the cost of inefficient fluctuations in international relative prices, which are also due to the fixed exchange rate.

With few exceptions, these conclusions are not to any important degree sensitive to the price-setting assumption (producer-currency pricing versus local-currency pricing) or the type of shocks (productivity shocks versus cost-push shocks). However, local-currency pricing and cost-push shocks—individually as well as jointly—tend to increase the likelihood that countries benefit from monetary unification. Compared with producer-currency pricing, local-currency pricing renders monetary unification more favorable because the benefit of exchange rate flexibility in the presence of nominal price rigidity—and therefore the cost of fixing the exchange rate—is considerably smaller. Under local-currency pricing, import prices no longer fluctuate one-to-one with the exchange rate but are as sticky as domestic prices. Therefore, exchange rate flexibility no longer facilitates the desirable adjustment in international relative goods prices. Compared with productivity shocks, cost-push shocks render monetary unification more favorable because the inherent benefit of monetary unions mentioned above is stronger under these circumstances. Cost-push shocks induce (possibly additional) tradeoffs for monetary policy in stabilizing different welfare-relevant variables. The bigger these tradeoffs are, the greater is the benefit of stabilizing private-sector expectations about future inflation.

1.1 Contribution to the Literature

The idea that monetary policy is an important OCA criterion has been touched upon in the literature at best only indirectly. There are two basic arguments: According to the "credibility" argument, a country that is unable to withstand the temptation to induce surprise inflations in a discretionary way suffers from both a higher level and a higher instability of inflation.[2] Joining or forming a monetary union can compensate for such a lack of commitment, thereby reducing the long-run level of inflation (Giavazzi and Pagano 1988; Alesina and Barro 2002; Chari, Dovis, and Kehoe 2020) and increasing the stability of inflation (Cook and Devereux 2016; Groll and Monacelli 2020).[3] According to the "competitive devaluations" argument, high and variable inflation arises when two countries with competing monetary policies try to strategically manipulate the real exchange rate or the terms of trade in their own favor. If the two countries form a monetary union, competitive devaluations are no longer possible and inflation is both lower (Cooley and Quadrini 2003) and more stable (Pappa 2004).

Without explicitly making the point, these contributions show en passant that monetary policy is an important OCA criterion. That is, whether countries are better off with flexible exchange rates or in a monetary union depends on whether their monetary authorities credibly commit to future policies (commitment versus discretion) and whether they coordinate their policies (coordination versus competition). In this paper, I broaden the perspective by looking

[2]In practice, there are a number of reasons to pursue such a policy. In times with a private or public debt overhang, monetary policy might let the inflation rate overshoot the inflation target for a prolonged period of time, with the aim of reducing the real debt burden and lowering borrowing costs. In times of high unemployment, this might seem attractive because it reduces real wages in the presence of fixed-term nominal wage contracts, thereby increasing the demand for labor.

[3]Note that the "currency union" in Alesina and Barro (2002) refers to a situation where a client country unilaterally adopts the currency of an anchor country—a situation also known as dollarization. Nevertheless, the benefit of eliminating an inflation bias also exists if the client and anchor country form a monetary union where the common monetary policy inherits the credibility of the anchor country. The "advantage of tying one's hands" in Giavazzi and Pagano (1988) follows the same logic, while referring to the former European Monetary System (1979–99).

through the lens of practical interest rate rules, thereby highlighting the importance of two different dimensions of monetary policy: the choice of the target variables that monetary policy responds to and the strength of the response to these variables. This enables me to describe the implications of a wide variety of interest rate policies reflecting the diversity of monetary policy in practice. It is important to realize that monetary policy represents an OCA criterion not only under these suboptimal interest rate rules but also under optimal monetary policy. This is shown by the studies mentioned above, which are all based on some form of optimal monetary policy.

The remainder of this paper is organized as follows. Section 2 briefly outlines the structure of the model. Section 3 shows how different interest rate policies lead to different welfare rankings between a monetary union and a flexible exchange rate regime. Section 4 shows how different interest rate policies change the nature of the traditional OCA criterion of trade openness. Section 5 concludes.

2. Model

The model I use is a standard two-country New Keynesian dynamic stochastic general equilibrium (DSGE) model, and thus I provide only a very brief description. The model features two currency regimes:

(i) A monetary union (MU) regime: Both countries share the same currency. A common monetary policy governs the common nominal interest rate.

(ii) A flexible exchange rate (FX) regime: Each country maintains its national currency and conducts its own, independent monetary policy. Nominal interest rates are country specific. The nominal exchange rate between the two currencies is flexible.

The FX version of the model, including the microfounded, quadratic welfare measure, is described in Corsetti, Dedola, and Leduc (2011). The MU version of the model is largely identical (see, e.g., Benigno 2004). The model economy features two countries of equal size (labeled H and F) with trade in consumption goods. The consumption baskets are allowed to differ among countries, so

purchasing power parity does not necessarily hold. International asset markets are complete, i.e., risk sharing is perfect across countries. Producers act in an environment of monopolistic competition. The only factor of production is labor, which is immobile between countries. The only rigidity is the nominal price rigidity in the spirit of Calvo (1983).

Under the FX regime, the baseline model assumes "producer-currency pricing." Prices are set in the currency of the producer's country. The price of imports expressed in domestic currency fluctuates one-to-one with the nominal exchange rate. Thus, the law of one price holds and exchange rate pass-through to import prices is complete. This implies that import prices are not sticky even though prices for domestically produced goods are sticky.

In order to check whether the main results are sensitive to the price-setting assumption, the case of "local-currency pricing" is also considered. Under local-currency pricing, prices are set in the currency of the destination market, i.e., in domestic currency if the good is sold domestically, and in foreign currency if the good is sold abroad. This implies that not only prices for domestically produced goods but also import prices are sticky. As a result, exchange rate pass-through is incomplete, and fluctuations in the nominal exchange rate lead to temporary deviations from the law of one price.

Under the MU regime, local-currency pricing is literally impossible, as both countries share one common currency. While other forms of price discrimination are clearly conceivable within a monetary union, modeling them is beyond the scope of this paper. Thus, the law of one price is assumed to always hold under the MU regime.[4]

2.1 Model Equations

The equations of the complete log-linearized model are shown below (for the full derivation, see appendixes A and B). Deviations of the logarithm of a variable X_t from its steady state are denoted by $\hat{X}_t$

[4]To the best of my knowledge, price discrimination within a monetary union has not yet been modeled in the New Keynesian open-economy macroeconomics literature. Interestingly, there is empirical evidence supporting the idea that the law of one price holds within a monetary union but not outside (see, e.g., Cavallo, Neiman, and Rigobon 2014).

Table 1. Variables

C_t, C_t^*	Consumption in Country H and F, Respectively
$Y_{H,t}, Y_{F,t}$	Output in Country H and F, Respectively
$\pi_{H,t}$	Producer Price Inflation in Country H in Country H's Currency
$\pi_{H,t}^*$	Producer Price Inflation in Country H in Country F's Currency
π_{Ft}	Producer Price Inflation in Country F in Country H's Currency
$\pi_{F,t}^*$	Producer Price Inflation in Country F in Country F's Currency
π_t, π_t^*	Consumer Price Inflation in Country H and F, Respectively
π_t^{MU}	Union-wide Inflation (Average of Country-Specific Inflation)
R_t, R_t^*	Nominal Interest Rate in Country H and F, Respectively
R_t^{MU}	Nominal Interest Rate in Monetary Union
T_t	Terms of Trade
S_t	Nominal Exchange Rate
Q_t	Real Exchange Rate
Δ_t	Deviation from the Law of One Price
$\zeta_{Y,t}, \zeta_{Y,t}^*$	Productivity Shock in Country H and F, Respectively
$\zeta_{C,t}, \zeta_{C,t}^*$	Consumption Preference Shock in Country H and F, Respectively
μ_t^H, μ_t^F	Cost-Push (or Markup) Shock in Country H and F, Respectively

Table 2. Parameters and Baseline Calibration

ρ	1/6	Inverse of Elasticity of Intertemporal Substitution in Consumption
β	0.99	Discount Factor
η	0.67	Inverse of Elasticity of Producing the Differentiated Good
ϵ_{wy}	0.5	Production Elasticity of Average Real Wage
γ	0.75	Labor Income Share
a	0.75	Home Bias/Degree of Trade Openness
σ	7.66	Elasticity of Substitution between Differentiated Goods within Countries
θ	2	Elasticity of Substitution between Goods across Countries
α	0.75	Probability of Not Being Able to Reset the Price

if prices are sticky and by $\tilde{X}_t^{fb}$ if prices are flexible and markups are neutralized (efficient allocation). The variables and parameters are defined in tables 1 and 2, respectively.

2.1.1 Sticky-Price Model under the FX Regime

Producer-Currency Pricing. Under sticky prices, the model equations for the FX regime and producer-currency pricing are given by

$$E_t\hat{C}_{t+1} = \hat{C}_t + \frac{1}{\rho}\left(\hat{R}_t - E_t\pi_{t+1} + E_t\hat{\zeta}_{C,t+1} - \hat{\zeta}_{C,t}\right) \tag{1}$$

$$\hat{Q}_t = \rho\left(\hat{C}_t - \hat{C}_t^*\right) + \left(\hat{\zeta}_{C,t}^* - \hat{\zeta}_{C,t}\right) \tag{2}$$

$$E_t\Delta\hat{S}_{t+1} = \hat{R}_t - \hat{R}_t^* \tag{3}$$

$$\hat{Q}_t = (2a-1)\hat{T}_t \tag{4}$$

$$\hat{Y}_{H,t} = 2a(1-a)\theta\hat{T}_t + a\hat{C}_t + (1-a)\hat{C}_t^* \tag{5}$$

$$\hat{Y}_{F,t} = -2a(1-a)\theta\hat{T}_t + (1-a)\hat{C}_t + a\hat{C}_t^* \tag{6}$$

$$\begin{aligned}\pi_{H,t} = {} & (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}_{H,t}^{fb}\right) - 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^H + \beta E_t\pi_{H,t+1}\end{aligned} \tag{7}$$

$$\begin{aligned}\pi_{F,t}^* = {} & (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}_{F,t}^{fb}\right) + 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^F + \beta E_t\pi_{F,t+1}^*\end{aligned} \tag{8}$$

$$\hat{T}_t = \hat{T}_{t-1} + \pi_{F,t}^* - \pi_{H,t} + \Delta\hat{S}_t \tag{9}$$

$$\pi_t = a\pi_{H,t} + (1-a)(\pi_{F,t}^* + \Delta\hat{S}_t) \tag{10}$$

$$\pi_t^* = (1-a)(\pi_{H,t} - \Delta\hat{S}_t) + a\pi_{F,t}^*, \tag{11}$$

where

$$k = \frac{(1-\alpha\beta)(1-\alpha)}{\alpha}\frac{1}{1+\sigma\eta}. \tag{12}$$

Monetary policy in each country can respond to some measure of inflation, to some measure of output, and to the nominal exchange rate, and it can engage in interest rate smoothing. The specific functional forms of the interest rate rules will be shown in section 3.

Local-Currency Pricing. Under sticky prices, the model equations for the FX regime and local-currency pricing are given by

$$E_t\hat{C}_{t+1} = \hat{C}_t + \frac{1}{\rho}\left(\hat{R}_t - E_t\pi_{t+1} + E_t\hat{\zeta}_{C,t+1} - \hat{\zeta}_{C,t}\right) \tag{13}$$

$$\hat{Q}_t = \rho\left(\hat{C}_t - \hat{C}_t^*\right) + \left(\hat{\zeta}_{C,t}^* - \hat{\zeta}_{C,t}\right) \tag{14}$$

$$E_t \Delta \hat{S}_{t+1} = \hat{R}_t - \hat{R}_t^* \tag{15}$$

$$\hat{Q}_t = (2a-1)\hat{T}_t + 2a\hat{\Delta}_t \tag{16}$$

$$\hat{Y}_{H,t} = 2a(1-a)\theta(\hat{T}_t + \hat{\Delta}_t) + a\hat{C}_t + (1-a)\hat{C}_t^* \tag{17}$$

$$\hat{Y}_{F,t} = -2a(1-a)\theta(\hat{T}_t + \hat{\Delta}_t) + (1-a)\hat{C}_t + a\hat{C}_t^* \tag{18}$$

$$\begin{aligned}\pi_{H,t} = {} & (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}_{H,t}^{fb}\right) \\ & - (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}_t^{fb} + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] \\ & + k\hat{\mu}_t^H + \beta E_t \pi_{H,t+1}\end{aligned} \tag{19}$$

$$\begin{aligned}\pi_{H,t}^* = {} & (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}_{H,t}^{fb}\right) \\ & - (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}_t^{fb} + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] - k\hat{\Delta}_t \\ & + k\hat{\mu}_t^H + \beta E_t \pi_{H,t+1}^*\end{aligned} \tag{20}$$

$$\begin{aligned}\pi_{F,t} = {} & (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}_{F,t}^{fb}\right) \\ & + (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}_t^{fb} + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] + k\hat{\Delta}_t \\ & + k\hat{\mu}_t^F + \beta E_t \pi_{F,t+1}\end{aligned} \tag{21}$$

$$\begin{aligned}\pi_{F,t}^* = {} & (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}_{F,t}^{fb}\right) \\ & + (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}_t^{fb} + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] \\ & + k\hat{\mu}_t^F + \beta E_t \pi_{F,t+1}^*\end{aligned} \tag{22}$$

$$\hat{T}_t = \hat{T}_{t-1} + \pi_{F,t} - \pi_{H,t}^* - \Delta\hat{S}_t \tag{23}$$

$$\pi_t = a\pi_{H,t} + (1-a)\pi_{F,t} \tag{24}$$

$$\pi_t^* = (1-a)\pi_{H,t}^* + a\pi_{F,t}^* \tag{25}$$

$$\hat{\Delta}_t = \hat{\Delta}_{t-1} + \Delta S_t + \pi_{H,t}^* - \pi_{H,t}. \tag{26}$$

2.1.2 Sticky-Price Model under the MU Regime

Under sticky prices, the model equations for the MU regime are given by

$$E_t\hat{C}_{t+1} = \hat{C}_t + \frac{1}{\rho}\left(\hat{R}_t^{MU} - E_t\pi_{t+1} + E_t\hat{\zeta}_{C,t+1} - \hat{\zeta}_{C,t}\right) \tag{27}$$

$$\hat{Q}_t = \rho\left(\hat{C}_t - \hat{C}_t^*\right) + \left(\hat{\zeta}_{C,t}^* - \hat{\zeta}_{C,t}\right) \tag{28}$$

$$\hat{Q}_t = (2a-1)\hat{T}_t \tag{29}$$

$$\hat{Y}_{H,t} = 2a(1-a)\theta\hat{T}_t + a\hat{C}_t + (1-a)\hat{C}_t^* \tag{30}$$

$$\hat{Y}_{F,t} = -2a(1-a)\theta\hat{T}_t + (1-a)\hat{C}_t + a\hat{C}_t^* \tag{31}$$

$$\begin{aligned}\pi_{H,t} = {} & (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}_{H,t}^{fb}\right) - 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^H + \beta E_t\pi_{H,t+1}\end{aligned} \tag{32}$$

$$\begin{aligned}\pi_{F,t}^* = {} & (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}_{F,t}^{fb}\right) + 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^F + \beta E_t\pi_{F,t+1}^*\end{aligned} \tag{33}$$

$$\hat{T}_t = \hat{T}_{t-1} + \pi_{F,t}^* - \pi_{H,t} \tag{34}$$

$$\pi_t = a\pi_{H,t} + (1-a)\pi_{F,t}^* \tag{35}$$

$$\pi_t^* = (1-a)\pi_{H,t} + a\pi_{F,t}^*. \tag{36}$$

The common monetary policy responds to union-wide variables, i.e., to cross-country averages. The specific functional forms of the interest rate rule will be shown in section 3.

Note that whether the common monetary policy responds to producer price inflation or consumer price inflation does not make a difference in this model, given that the two countries are of equal size. Using equations (35) and (36), it is straightforward to show that the average of consumer price inflation rates is equal to the average of producer price inflation rates:

$$\frac{\pi_t + \pi_t^*}{2} = \frac{\pi_{H,t} + \pi_{F,t}^*}{2} \equiv \pi_t^{MU}. \tag{37}$$

2.1.3 Efficient Allocation

The following equations describe the first-best (fb) or efficient allocation, where prices are fully flexible, where the law of one price holds, and where markups are neutralized at all times with an appropriate subsidy ($\mu_t^i = 0$). This efficient allocation provides a useful benchmark for assessing the welfare implications of the two currency regimes.

The efficient output in each country is given by

$$(\rho+\eta)\tilde{Y}_{H,t}^{fb} = 2a(1-a)(\rho\theta-1)\tilde{T}_t^{fb} - (1-a)\left(\hat{\zeta}_{C,t} - \hat{\zeta}_{C,t}^*\right) + \hat{\zeta}_{C,t} + \eta\hat{\zeta}_{Y,t} \tag{38}$$

$$(\rho+\eta)\tilde{Y}_{F,t}^{fb} = -2a(1-a)(\rho\theta-1)\tilde{T}_t^{fb} + (1-a)\left(\hat{\zeta}_{C,t} - \hat{\zeta}_{C,t}^*\right) + \hat{\zeta}_{C,t}^* + \eta\hat{\zeta}_{Y,t}^*. \tag{39}$$

The efficient terms of trade can be written as

$$[4a(1-a)\rho\theta + (2a-1)^2]\tilde{T}_t^{fb} = \rho\left(\tilde{Y}_{H,t}^{fb} - \tilde{Y}_{F,t}^{fb}\right) - (2a-1)\left(\hat{\zeta}_{C,t} - \hat{\zeta}_{C,t}^*\right). \tag{40}$$

2.2 Model Description

Producer-Currency Pricing. Consumption growth is described by standard Euler equations, which are given by equations (1) and (27) in the case of country H. The difference between these two Euler equations is that the nominal interest rate is country specific under the FX regime, whereas it is common to both countries under the MU regime. The risk-sharing condition, which describes the link between consumption across countries, is identical across regimes and it is given by (2) and (28), respectively. Purchasing power parity does not hold at all times, i.e., the real exchange rate is not constant, unless consumption and consumption preference shocks are perfectly correlated across countries. Under the FX regime, perfect risk sharing implies that the uncovered interest parity (3) holds, i.e., the expected change in the nominal exchange rate corresponds

to the interest rate differential across countries.[5] This equation is obsolete under the MU regime because both countries share the same currency and a common nominal interest rate.

The link between the real exchange rate and the terms of trade is described by equations (4) and (29), respectively. Accordingly, the correlation between the real exchange rate and the terms of trade can be positive, zero, or negative, depending on the degree of trade openness between the two countries. Aggregate demand in each country depends on consumption in both countries and the terms of trade and is given by equations (5), (6), (30), and (31), respectively. The country-specific New Keynesian Phillips curves are also identical across regimes and they are given by (7), (8), (32), and (33), respectively. In contrast to a closed-economy framework, not only the output gap but also the terms-of-trade gap (the difference between the sticky price and the efficient terms of trade) affect producer price inflation. I follow much of the related literature in modeling cost-push shocks in an ad hoc way as exogenous fluctuations in the markup μ_t induced by time-varying taxes.

The terms-of-trade identity is given by equation (9) under the FX regime and by equation (34) under the MU regime, which differ due to the presence of the nominal exchange rate in the former. Equations (10), (11), (35), and (36) describe the relationship between the consumer price inflation rate and the producer price inflation rates in each country. Likewise, these equations only differ across regimes in terms of the presence of the nominal exchange rate.

Under flexible prices, monetary policy is neutral and real variables are driven only by productivity shocks and consumption preference shocks. Thus, the efficient allocation, which is given by equations (38) through (40), is the same under both currency regimes.

Local-Currency Pricing. The Euler consumption equation (13), the risk-sharing condition (14), and the uncovered interest parity condition (15) are identical to the case of producer-currency pricing. The real exchange rate is still linked to the terms of trade,

[5]Combining the Euler consumption equation for country H, the risk-sharing condition, and the uncovered interest parity condition yields the Euler consumption equation for country F, which is therefore redundant. Alternatively, the model can be specified by including both country-specific Euler consumption equations and the risk-sharing condition, while omitting the uncovered interest parity condition.

but it is now also linked to the deviation from the law of one price (equation (16)). Since the countries are assumed to be symmetric, the deviation from the law of one price is identical across countries ($\hat{\Delta}_{H,t} = \hat{\Delta}_{F,t} = \hat{\Delta}_t$). The aggregate demand equations (17) and (18) as well as the four New Keynesian Phillips curves (19) through (22) contain the deviation from the law of one price as well. The terms-of-trade identity (23) and the definitions of the CPI inflation rates (24) and (25) are different from the case of producer-currency pricing, since the law of one price does not hold under local-currency pricing. Finally, equation (26) defines the deviation from the law of one price, expressed in first differences.

2.3 Welfare Loss Function

The welfare analysis follows the logic of the familiar linear-quadratic approach, where the log-linear model equations are used to evaluate a quadratic welfare loss measure (Woodford 2003). The joint welfare loss function is given by the discounted value of a weighted average across countries of the average utility flow of agents using a second-order Taylor-series expansion.[6] It is assumed that the distortion induced by monopolistic competition is offset by an appropriate subsidy, thereby ensuring efficiency in the steady state.

Producer-Currency Pricing. The welfare loss function in the case of producer-currency pricing is given by (see Corsetti, Dedola, and Leduc 2011):

$$\begin{aligned} W_t = -\frac{1}{2}\Bigg(& (\rho+\eta)\operatorname{var}\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) + (\rho+\eta)\operatorname{var}\left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right) \\ & - \frac{2a(1-a)(\rho\theta-1)\rho}{4a(1-a)\rho\theta + (2a-1)^2}\operatorname{var}\left[\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) - \left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right)\right] \\ & + \frac{\sigma}{k}\left[\operatorname{var}\pi_{H,t} + \operatorname{var}\pi^*_{F,t}\right]\Bigg) + t.i.p. + O(\|\xi\|^3). \end{aligned} \tag{41}$$

[6]Computing country-specific welfare would complicate the calculations significantly because more accurate approximations of the nonlinear model equations would be necessary (Benigno and Woodford 2005), which is beyond the scope of this paper. As long as the countries are symmetric, a gain in joint welfare always implies a gain for both countries. There is only one case where asymmetric countries are considered (section 3.3).

The weights in front of each component of the welfare loss function are functions of the deep parameters of the model. The term $t.i.p.$ contains all the terms that are independent of monetary policy and the currency regime. The term $O(\|\xi\|^3)$ contains third- and higher-order terms, which can be neglected provided that the model equations are log-linear, i.e., first-order approximations of the nonlinear equilibrium conditions.

As in the closed economy, the welfare loss depends on the producer price inflation rate and the output gap. In the open economy, the welfare loss also depends on the output gap differential across countries. If the output gap differential is different from zero, the allocation of production *across* countries is inefficient. Importantly, under producer-currency pricing, the output gap differential and the terms-of-trade gap are two sides of the same coin. To see this, combine equation (40) with its analogous sticky-price counterpart to obtain

$$\begin{aligned}&\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) - \left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right)\\ &\quad = \frac{4a(1-a)\rho\theta + (2a-1)^2}{\rho}\left(\hat{T}_t - \tilde{T}^{fb}_t\right). \qquad (42)\end{aligned}$$

Thus, stabilizing the output gap differential automatically stabilizes the terms-of-trade gap, and vice versa. And the welfare loss function above can be expressed in terms of the terms-of-trade gap instead of the output gap differential, which I will make use of in the analysis.

Local-Currency Pricing. The welfare loss function in the case of local-currency pricing is given by (see Corsetti, Dedola, and Leduc 2011):

$$\begin{aligned}W_t = -\frac{1}{2}\Bigg(&(\rho+\eta)\operatorname{var}\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) + (\rho+\eta)\operatorname{var}\left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right)\\ &- \frac{2a(1-a)(\rho\theta-1)\rho}{4a(1-a)\rho\theta + (2a-1)^2}\operatorname{var}\left[\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) - \left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right)\right]\\ &+ \frac{2a(1-a)\theta}{4a(1-a)\rho\theta + (2a-1)^2}\operatorname{var}\hat{\Delta}_t\\ &+ \frac{\sigma}{k}\big[a\operatorname{var}\pi_{H,t} + (1-a)\operatorname{var}\pi^*_{H,t}\end{aligned}$$

$$+ a \operatorname{var} \pi^*_{F,t} + (1-a) \operatorname{var} \pi_{F,t}\Big]\Big)$$
$$+ t.i.p. + O(\|\xi\|^3). \tag{43}$$

Compared with the case of producer-currency pricing, the welfare loss function under local-currency pricing contains additional terms. First, it depends on the deviation from the law one price. Second, it depends on all four producer price inflation rates.[7] Importantly, under local-currency pricing, the output gap differential and the terms-of-trade gap are no longer two sides of the same coin. Following the same steps as before yields

$$\begin{aligned}\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) - \left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right) \\ = \frac{4a(1-a)\rho\theta + (2a-1)^2}{\rho}\left(\hat{T}_t - \tilde{T}^{fb}_t\right) \\ + \frac{4a(1-a)\rho\theta + 2a(2a-1)}{\rho}\hat{\Delta}_t.\end{aligned} \tag{44}$$

Thus, stabilizing the output gap differential does not automatically stabilize the terms-of-trade gap, and vice versa, because of potential deviations from the law of one price.

2.4 Calibration

Unless stated otherwise, the parameters of the model are calibrated to the values displayed in table 2 (see also Benigno 2004). For the sake of simplicity, the two countries are assumed to be symmetric. A value of 0.99 for the discount factor β implies a steady-state real interest rate of around 4.1 percent annually. A value of 7.66 for the elasticity of substitution between differentiated goods σ implies a steady-state markup of prices over marginal costs of 15 percent. A

[7] Recall that inflation rates are relevant for welfare losses because they imply inefficient price dispersion in the presence of staggered price setting. Thus, the reason the welfare loss function under *producer*-currency pricing only contains two inflation rates is that the dispersion of prices of, e.g., domestically produced goods is identical in domestic and foreign currency (Engel 2011). It is not because the inflation rate for one good in different currencies is identical, which generally it is not.

value of 0.75 for the probability of not being able to reset the price α implies an average duration of price contracts of four quarters.

The degree of trade openness a is calibrated to 0.75, which corresponds to a steady-state share of home-produced goods in the consumption basket of 75 percent in each country (i.e., a home bias in consumption) and a steady-state trade-to-GDP ratio of 50 percent.[8] This roughly equals the average trade-to-GDP ratio across OECD countries. Following Rotemberg and Woodford (1998) and Benigno (2004), the inverse of the elasticity of producing the differentiated good η is calculated as

$$\eta = \epsilon_{wy} - \rho + \frac{1-\gamma}{\gamma}, \tag{45}$$

where ϵ_{wy} denotes the elasticity of the average real wage with respect to production and γ denotes the labor income share.

With the exception of the exchange rate coefficient ϕ_S, all interest rate rule coefficients are assumed to be identical across countries and regimes. Finally, the persistence of shocks is set to 0.9 in each country, and the cross-country correlation of shocks is zero.

3. Monetary Policy as an OCA Criterion

In the following, I use the theoretical model described in the previous section to show that the conduct of monetary policy is a critical criterion for the welfare implications of monetary unification. The conduct of monetary policy can differ with respect to the coefficients in the interest rate rules that determine the response of monetary policy to inflation (section 3.1), to output (section 3.2), to the nominal exchange rate (section 3.3), and to past realizations of the interest rate (section 3.4). In addition, the conduct of monetary policy can differ with respect to the target variables themselves. Monetary policy can respond to producer price inflation (henceforth PPI inflation targeting) or consumer price inflation (henceforth CPI inflation targeting), and it can respond to output (deviation from steady state) or the output gap (deviation from efficient allocation).

The baseline results are shown for producer-currency pricing and productivity shocks. In addition, I discuss the cases of local-currency

[8]The steady-state trade-to-GDP ratio in percent is given by $2(1-a) \times 100$.

pricing and cost-push shocks to stress that the results are not to any important degree sensitive to these modeling choices.

3.1 Response to Inflation

Under the FX regime and PPI inflation targeting, the interest rate rules for both countries are given by

$$\hat{R}_t = \phi_\pi \pi_{H,t} \tag{46}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_{F,t}. \tag{47}$$

Under CPI inflation targeting, they take the following form:

$$\hat{R}_t = \phi_\pi \pi_t \tag{48}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_t. \tag{49}$$

Under the MU regime, the interest rate rule of the common monetary policy is the same under PPI and CPI inflation targeting (recall equation (37)):

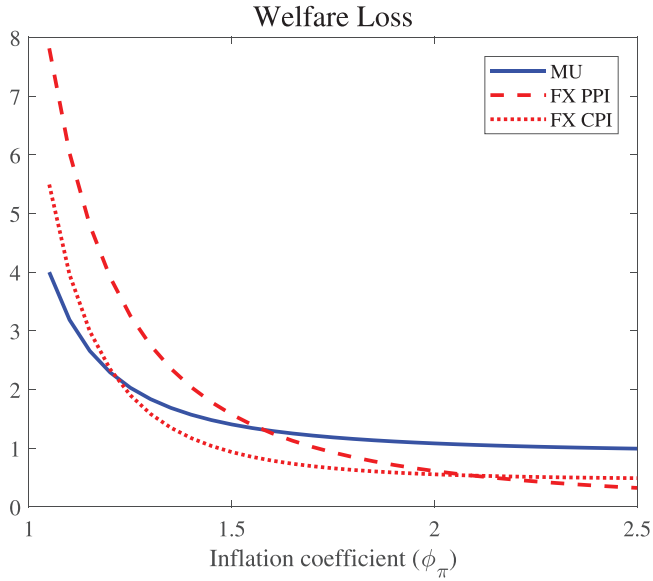
$$\hat{R}^{MU}_t = \phi_\pi \pi^{MU}_t. \tag{50}$$

Producer-Currency Pricing and Productivity Shocks. The aggressiveness of monetary policy in its response to inflation has a determining influence on whether countries are better off under the MU regime or under the FX regime (figure 1). If the response to inflation is relatively modest (i.e., low values of ϕ_π), the two countries are better off under the MU regime. If the response to inflation is relatively strong, the two countries are better off under the FX regime. The threshold value of ϕ_π beyond which the FX regime becomes superior depends on the measure of inflation monetary policy responds to. Under CPI inflation targeting, the threshold value for ϕ_π is lower than under PPI inflation targeting.

The welfare ranking between the MU and the FX regime is driven by the inflation component, which exhibits the same pattern with respect to ϕ_π as the welfare loss (figure 2, lower right panel).[9] This is

[9] Although the welfare loss depends on the output gap and the PPI inflation rate of both countries, figure 2 shows only one of each because the variances are identical due to the assumption of symmetric countries.

Figure 1. Welfare Loss as a Function of the Inflation Coefficient (ϕ_π) under Producer-Currency Pricing and Productivity Shocks

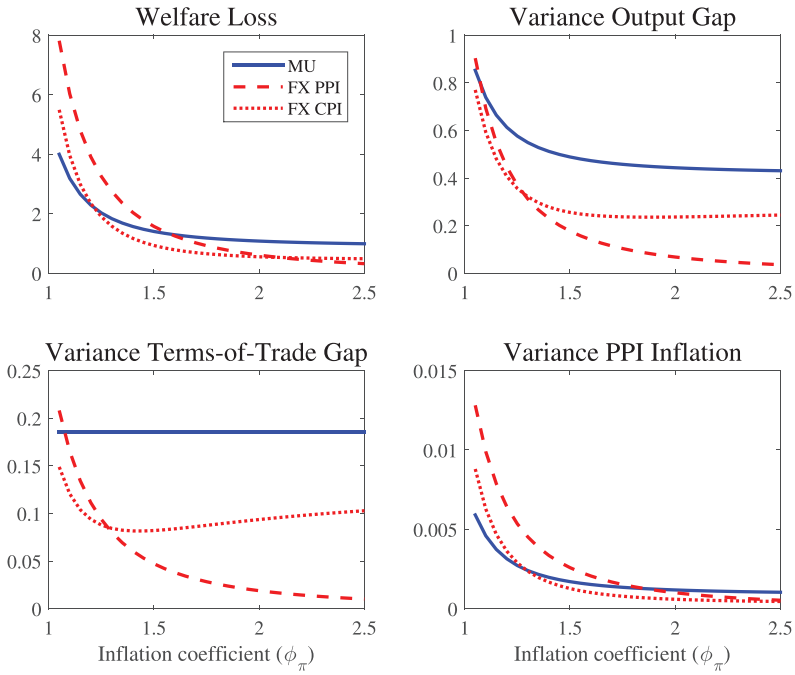


because agents attach by far the highest weight to inflation, which is traditionally the case in microfounded welfare measures derived from New Keynesian models.[10] Accordingly, the cost of a higher variance of the output gap and of the terms-of-trade gap—or, equivalently, of the output gap differential (recall equation (42))—under the MU regime (figure 2, upper right and lower left panel) can be outweighed by the benefit of a lower variance of PPI inflation. This is the case for low values of ϕ_π, i.e., a relatively modest response of monetary policy to inflation.

The reason why the two countries are better off under the FX regime for a sufficiently strong response of monetary policy to inflation is predominantly related to the effectiveness of monetary policy.

[10]Under the baseline calibration, the coefficients in front of the inflation rate, the output gap, and the terms-of-trade gap in the welfare loss function are 555.98, 0.83, and 0.75, respectively.

Figure 2. Welfare Loss and Variances of Welfare-Relevant Variables as a Function of the Inflation Coefficient (ϕ_π) under Producer-Currency Pricing and Productivity Shocks



This becomes clear by comparing the number of policy instruments with the number of welfare-relevant distortions in the economy.

Under the FX regime, there are as many policy instruments as distortions in the two-country world (four). The distortions are due to monopolistic competition and to sticky prices in each country.[11] The distortion due to monopolistic competition induces an inefficiently low level of aggregate output. This distortion can be eliminated by an appropriate subsidy in each country. The distortion due to sticky prices induces inefficient markup fluctuations, which lead

[11]Both distortions are common to the closed-economy framework (see, e.g., Woodford 2003 for details).

to inefficiently low or high levels of aggregate output, and an inefficient dispersion of prices in the presence of inflation, which causes an inefficient dispersion of output across the producers of differentiated goods *within* each country. This distortion can be mitigated or even eliminated by monetary policy in each country by using the nominal interest rate to reduce the fluctuations of inflation around zero as far as possible.

By contrast, under the MU regime, there are fewer policy instruments (three) than distortions (five) in the two-country world. First, monetary policy sets the nominal interest rate for both countries and thus it can no longer target inflation in each country separately, thereby losing one policy instrument. Second, the combination of the fixed nominal exchange rate with sticky prices induces an additional distortion, namely an intrinsic inertia in the terms of trade (Benigno 2004; Pappa 2004; Groll and Monacelli 2020).[12] This causes an inefficient dispersion of aggregate output *across* countries.

Given that there are as many policy instruments as distortions under the FX regime but fewer policy instruments than distortions under the MU regime, monetary policy is more effective under the FX regime, which shows up clearly in figure 2. The "leverage" of monetary policy is higher under the FX regime than under the MU regime in the sense that a given increase in the aggressiveness of monetary policy toward inflation (measured by ϕ_π) leads to a larger reduction in the variance of each welfare-relevant variable. In fact, under PPI inflation targeting, monetary policy can reduce the variances of all welfare-relevant variables to zero ($\phi_\pi \to \infty$). This is impossible under the MU regime.[13]

[12]Intrinsic inertia is defined as follows: Consider a one-off (i.e., nonpersistent) productivity shock in one country. Under the MU regime, several periods are required before the terms of trade return to the steady state after the shock has vanished. The terms of trade are said to be intrinsically persistent or inertial. Under the FX regime, the terms of trade return to the steady state immediately after the shock has vanished. In this case, the terms of trade are not intrinsically inertial.

[13]See Groll (2013) for the analytical proof in the special case where $a = 1/2$ and $\theta = 1$. The proof in the case of no restrictions on a and θ is completely analogous.

The limitations of monetary policy under the MU regime apply in particular to the terms-of-trade gap or, equivalently, to the output gap differential (figure 2, lower left panel). Monetary policy has no effect whatsoever on the terms of trade and thus on the terms-of-trade gap. Since both countries face the same nominal interest rate, any interest rate adjustment by the common monetary policy has the same initial effect on both countries. If the degree of price stickiness is identical across the two countries, an interest rate adjustment propagates through both economies in exactly the same way, and the influence of monetary policy on the terms of trade is zero. If the degree of price stickiness were not identical across the two countries, the influence of monetary policy on the terms of trade would not be zero, but would still be very small.

Despite those limitations of monetary policy, countries can be better off under the MU regime, as is the case for a relatively modest response of monetary policy to inflation. Paradoxically, the intrinsic inertia in the terms of trade due to the fixed exchange rate can also be beneficial, as is explained in detail in Groll and Monacelli (2020). In short, the inertia in the terms of trade has the advantage of stabilizing private-sector expectations about future inflation and thereby stabilizing actual inflation. This can overcompensate for the cost of inefficient terms-of-trade fluctuations, which are also induced by the fixed exchange rate. I will refer to this "inherent benefit of monetary unions" a number of times throughout the paper.

Robustness. Under either local-currency pricing or cost-push shocks, it continues to hold that the countries are better off under the MU regime if monetary policy responds to inflation modestly (see appendix C, figure C.1). However, the threshold value of ϕ_π beyond which the FX regime becomes superior is generally higher compared with the case of producer-currency pricing or productivity shocks. Thus, the MU regime is more likely to be superior. If local-currency pricing and cost-push shocks concur, the MU regime is superior irrespective of the aggressiveness of monetary policy toward inflation.

Compared with productivity shocks, cost-push shocks render the MU regime more favorable because the inherent benefit of monetary unions mentioned above is stronger under these circumstances. Cost-push shocks induce (possibly additional) tradeoffs for monetary policy in stabilizing different welfare-relevant variables. The bigger

these tradeoffs are, the greater is the benefit of stabilizing private-sector expectations about future inflation. This benefit is inherent to the MU regime due to the fixed exchange rate.[14]

Compared with producer-currency pricing, local-currency pricing renders the MU regime more favorable because the benefit of exchange rate flexibility in the presence of nominal price rigidity—and therefore the cost of fixing the exchange rate—is considerably smaller. Under local-currency pricing, import prices no longer fluctuate one-to-one with the exchange rate but are as sticky as domestic prices. Therefore, exchange rate flexibility no longer facilitates the desirable adjustment in international relative prices of goods ($\hat{T}_t + \hat{\Delta}_t$) in response to country-specific shocks (Devereux and Engel 2003; Corsetti, Dedola, and Leduc 2011; Engel 2011). There are more distortions than policy instruments, namely two sticky prices versus one interest rate within each country. As a result, monetary policy is less effective under local-currency pricing. Nevertheless, a case for flexible exchange rates remains even if there is no expenditure-switching effect of the exchange rate: Exchange rate flexibility facilitates the desirable adjustment in the real exchange rate ($\hat{Q}_t$), accommodating the efficient response of aggregate consumption across countries (Duarte and Obstfeld 2008). This explains why countries can be better off under the FX regime even under local-currency pricing.

3.2 Response to Output

In this subsection, monetary policy responds not only to inflation but also to output (deviation from steady state) or to the output gap (deviation from efficient allocation). Under the FX regime, if monetary policy targets output, the interest rate rules for both countries are given by

$$\hat{R}_t = \phi_\pi \pi_{H,t} + \phi_Y \hat{Y}_{H,t} \tag{51}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_{F,t} + \phi_Y \hat{Y}_{F,t}. \tag{52}$$

[14]The logic is completely analogous to the gains of optimal monetary policy under commitment. These gains also operate through expectations and are increasing in the severity of the tradeoffs faced by monetary policy (Woodford 2003).

If monetary policy targets the output gap, they take the following form:

$$\hat{R}_t = \phi_\pi \pi_{H,t} + \phi_Y \left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t} \right) \tag{53}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_{F,t} + \phi_Y \left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t} \right). \tag{54}$$

Under the MU regime, if the common monetary policy responds to output, the interest rate rule is given by

$$\hat{R}^{MU}_t = \phi_\pi \pi^{MU}_t + \phi_Y \frac{\hat{Y}_{H,t} + \hat{Y}_{F,t}}{2}. \tag{55}$$

If it responds to the output gap, it is given by

$$\hat{R}^{MU}_t = \phi_\pi \pi^{MU}_t + \phi_Y \frac{\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t} \right) + \left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t} \right)}{2}. \tag{56}$$

In all of these cases, the inflation coefficient ϕ_π is set to 1.5. As the difference between PPI and CPI inflation targeting is very small in this context, only the results under PPI inflation targeting are reported.

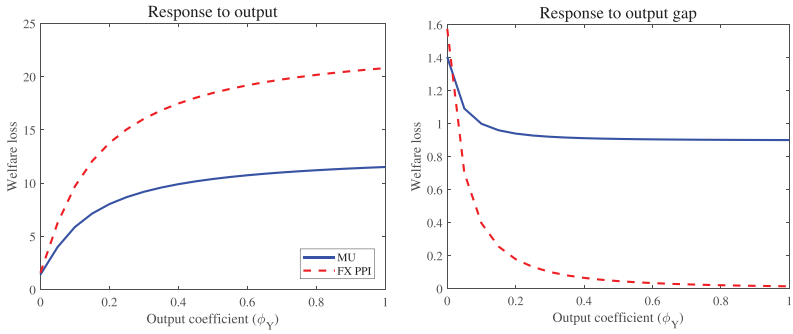
Producer-Currency Pricing and Productivity Shocks. Whether countries are better off under the MU regime or under the FX regime depends crucially on whether monetary policy responds to output (deviation from steady state) or the output gap (deviation from efficient allocation). If monetary policy responds to output, the two countries are better off under the MU regime (figure 3, left panel). By contrast, if monetary policy responds to the output gap, the two countries are better off under the FX regime (figure 3, right panel). As before, the driving factor is the inflation component.[15]

The key to understanding these results is the role played by the nominal exchange rate in stabilizing the terms-of-trade gap. Combining equations (38) to (40) and focusing on productivity shocks in country H yields the following relationship between the efficient terms of trade $\tilde{T}^{fb}_t$ and the productivity shock $\hat{\zeta}_{Y,t}$:

$$\tilde{T}^{fb}_t = \frac{\rho\eta}{4a(1-a)\rho(1+\eta\theta) + (\rho+\eta)(2a-1)^2} \hat{\zeta}_{Y,t}. \tag{57}$$

[15]Not shown, but available upon request.

Figure 3. Welfare Loss as a Function of the Output Coefficient (ϕ_Y) under Producer-Currency Pricing and Productivity Shocks



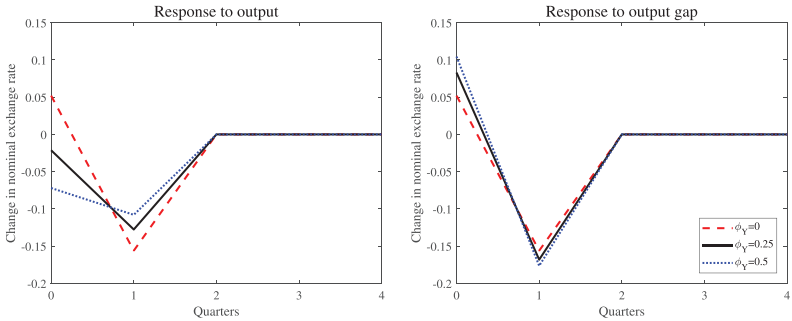
Notes: Left panel: Response to output ($\hat{Y}_t$). Right panel: Response to output gap ($\hat{Y}_t - \tilde{Y}_t^{fb}$).

The term in front of the productivity shock is unambiguously positive. Accordingly, the terms of trade would increase in response to a positive productivity shock in country H if prices were perfectly flexible. However, because prices are sticky, the actual increase in the terms of trade is smaller. In these circumstances, an increase in the nominal exchange rate would help to close the gap between the actual response of the terms of trade and its efficient counterpart, thereby stabilizing the terms-of-trade gap and reducing the welfare loss. But whether the nominal exchange rate stabilizes or destabilizes the terms-of-trade gap depends crucially on whether monetary policy responds to output or the output gap. This is because the nominal exchange rate is directly linked to the interest rates governed by monetary policy via the uncovered interest parity condition (3).

If monetary policy responds to neither output nor the output gap ($\phi_Y = 0$), the impact response of the nominal exchange rate to a positive productivity shock in country H is positive, i.e., country H's currency depreciates (figure 4).[16] Thus, the nominal exchange

[16]The degree of price stickiness was set to a low value ($\alpha = 0.2$) to ensure that the differences in the impulse responses are clearly visible. The differences for higher degrees of price stickiness are smaller but qualitatively the same.

Figure 4. Impulse Response of the Change in the Nominal Exchange Rate ($\Delta\hat{S}_t$) to a Positive One-Off Productivity Shock in Country H for Three Different Values of the Output Coefficient (ϕ_Y), with $\alpha = 0.2$, under Producer-Currency Pricing



Notes: Left panel: Response to output ($\hat{Y}_t$). Right panel: Response to output gap ($\hat{Y}_t - \tilde{Y}_t^{fb}$).

rate pushes the sticky-price terms of trade in the same direction as the efficient terms of trade, thereby stabilizing the terms-of-trade gap to some extent. If monetary policy responds to the output gap, the positive impact response of the nominal exchange rate becomes greater as ϕ_Y increases (figure 4, right panel). The stabilizing effect increases accordingly, further reducing the terms-of-trade gap and reducing the welfare loss. Since this stabilizing mechanism is absent under the MU regime, the countries are better off under the FX regime.

By contrast, as monetary policy starts to respond to *output*, the impact response of the nominal exchange rate first becomes smaller and then negative for already very small values of ϕ_Y (figure 4, left panel). A negative impact response means that the nominal exchange rate *destabilizes* the terms-of-trade gap by pushing the sticky-price terms of trade away from the efficient terms of trade. As a result, not only the terms-of-trade gap but also the output gap and the PPI inflation rate are destabilized, thereby increasing the welfare loss. Under these circumstances, the countries are better off with a fixed exchange rate.

Importantly, the nominal exchange rate amplifies a detrimental effect that is already present; it does not cause the detrimental effect. In a closed economy, a response of monetary policy to output is also detrimental to welfare (see, e.g., Galí 2015, chapter 4.4). It is not the deviation of output from the steady state that is welfare relevant; it is the deviation from the efficient counterpart (output gap). A positive productivity shock in country H induces an increase in output but a decrease in the output gap, because the increase in output is lower than the increase in efficient output. A welfare-oriented response of monetary policy would require a reduction in the interest rate due to the negative output gap. Instead, monetary policy raises the interest rate due to the rise in output.

For these reasons, a response of monetary policy to output is detrimental under both the FX regime and the MU regime (in figure 3, left panel, the welfare loss increases in ϕ_Y under both regimes). However, the detrimental effect is larger under the FX regime due to the amplification by the nominal exchange rate. As described above, monetary policy is more effective under the FX regime than under the MU regime in terms of macroeconomic stabilization because of the flexibility of the nominal exchange rate. The flipside of this is that monetary policy can do more harm when it is not conducted properly. Essentially, the nominal exchange rate does not compensate for monetary policy mistakes; it reinforces them. In this sense, the MU regime provides a protective mechanism against monetary policy mistakes.

Robustness. Under local-currency pricing, it continues to hold that monetary unification is beneficial when monetary policy responds to output, and costly when it responds to the output gap (see appendix C, figure C.2). However, the welfare loss differences between the two currency regimes are smaller than under producer-currency pricing. This is because exchange rate flexibility is less beneficial under local-currency pricing due to the missing effect on international relative goods prices, which reduces the effectiveness of monetary policy under the FX regime (see above).

Under cost-push shocks, the situation is a little different than under productivity shocks. Both the response to output and the response to the output gap are detrimental to welfare, and both policies render the MU regime superior to the FX regime (see appendix C, figure C.3). This is primarily due to the inherent benefit of

monetary unions, which is much stronger under cost-push shocks (see above).

As in a closed economy, there is no difference between targeting output and targeting the output gap because the two variables are identical under cost-push shocks (the efficient allocation is unaffected). Cost-push shocks move output/the output gap and inflation in opposite directions. Given this tradeoff, responding more aggressively to output/the output gap automatically reduces the response to inflation. As a result, the output gap becomes more stable, but inflation becomes less stable. This reduces welfare, as agents attach a higher weight to inflation. For this reason, a response of monetary policy to output/the output gap is detrimental to welfare in the presence of cost-push shocks.

This continues to hold in the open economy under both the FX and MU regime. However, while the stabilizing effect on the output gap—and now in addition on the output gap differential—is smaller under the MU regime, the destabilizing effect on inflation is also smaller under the MU regime. Both effects are due to the fixed exchange rate. While hampering the stabilization of output gaps due to inefficient fluctuations of international relative prices, the fixed exchange rate has the advantage of stabilizing private-sector expectations about future inflation and thereby actual inflation. Due to the higher weight of inflation stability, the MU regime turns out to be superior in terms of welfare if monetary policy targets output/the output gap in the presence of cost-push shocks.

3.3 Response to Nominal Exchange Rate

In this subsection, monetary policy responds to inflation and the nominal exchange rate.[17] I distinguish between unilateral exchange rate targeting, where only one of the two countries responds to the exchange rate, and bilateral exchange rate targeting, where both countries respond to the exchange rate symmetrically. Under unilateral exchange rate targeting and PPI inflation targeting, the interest rate rules for both countries are given by

[17]See, e.g., Calvo and Reinhart (2002) for empirical estimates on the number of countries that target the exchange rate.

$$\hat{R}_t = \phi_\pi \pi_{H,t} + \frac{\phi_S}{1-\phi_S}\hat{S}_t \tag{58}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_{F,t}. \tag{59}$$

Under CPI inflation targeting, they are given by

$$\hat{R}_t = \phi_\pi \pi_t + \frac{\phi_S}{1-\phi_S}\hat{S}_t \tag{60}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_t. \tag{61}$$

Under bilateral exchange rate targeting and PPI inflation targeting, the interest rate rules for both countries are given by

$$\hat{R}_t = \phi_\pi \pi_{H,t} + \frac{\phi_S}{1-\phi_S}\hat{S}_t \tag{62}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_{F,t} - \frac{\phi_S}{1-\phi_S}\hat{S}_t. \tag{63}$$

Under CPI inflation targeting, they are given by

$$\hat{R}_t = \phi_\pi \pi_t + \frac{\phi_S}{1-\phi_S}\hat{S}_t \tag{64}$$

$$\hat{R}^*_t = \phi_\pi \pi^*_t - \frac{\phi_S}{1-\phi_S}\hat{S}_t. \tag{65}$$

The coefficient $\phi_S \in [0,1)$ governs the strength of the response to the exchange rate. It ranges from a regime of full exchange rate flexibility ($\phi_S = 0$) to a fixed exchange rate regime ($\phi_S \to 1$) with hybrid regimes in between (Galí and Monacelli 2016).

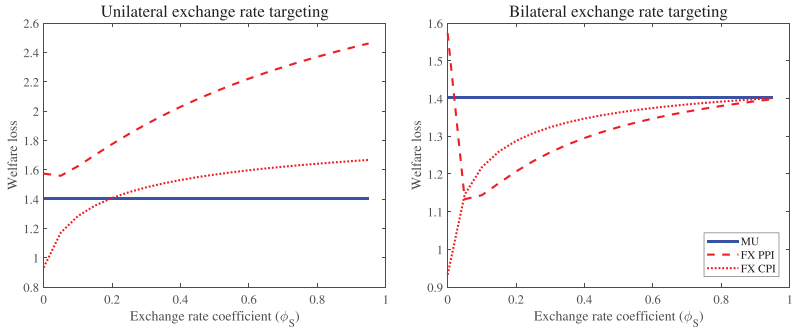
Under the MU regime, the interest rate rule is given by

$$\hat{R}^{MU}_t = \phi_\pi \pi^{MU}_t. \tag{66}$$

In all of these cases, the inflation coefficient ϕ_π is set to 1.5.

Producer-Currency Pricing and Productivity Shocks. Whether countries are better off under the MU regime or under the FX regime is not only sensitive to the degree to which countries respond to the nominal exchange rate but also, and more importantly, to whether the exchange rate targeting regime is carried out

Figure 5. Welfare Loss as a Function of the Exchange Rate Coefficient (ϕ_S) under Producer-Currency Pricing and Productivity Shocks



Notes: Left panel: Unilateral exchange rate targeting. Right panel: Bilateral exchange rate targeting.

unilaterally or bilaterally.[18] Under unilateral exchange rate targeting, the countries are generally better off under the MU regime (figure 5, left panel).[19] An exception is the case where monetary policy targets CPI inflation and responds to the exchange rate only very modestly. By contrast, under bilateral exchange rate targeting, the countries are generally worse off under the MU regime (figure 5, right panel). Here, the difference between CPI and PPI inflation targeting is small.

The principal reason for the different welfare implications of the unilateral and the bilateral exchange rate targeting regime vis-à-vis the MU regime are coordination gains. Consider the limiting case of a fixed exchange rate ($\phi_S \to 1$). Although the exchange rate is fixed under both unilateral and bilateral exchange rate targeting

[18]Both for simplicity and comparability with other sections, I continue to use the term "FX regime," although, clearly, targeting the exchange rate does not implement a regine in which the nominal exchange rate is perfectly flexible.

[19]In this particular case, it is possible that one country suffers a welfare loss, which is overcompensated by the other country's welfare gain. In all other welfare comparisons in this paper, a gain in joint welfare always implies a gain for both countries, due to the symmetry in country characteristics as well as in interest rate rules.

as well as under the MU regime, only the bilateral fixed exchange rate regime yields the same welfare as the MU regime. These two regimes are in fact identical in every respect. This is because both regimes implement the fixed exchange rate in a coordinated way. The MU regime represents a coordinated fixed exchange rate regime by construction. The bilateral fixed exchange rate regime implies coordination because both countries respond to the exchange rate symmetrically.

By contrast, under a unilateral fixed exchange rate regime (one-sided peg), only one of the two countries ensures that the exchange rate is fixed, while the other country can choose its interest rate policy independently. Since fixing the exchange rate requires the country-specific interest rates to be perfectly aligned at all times, the pegging country must always follow the other country's interest rate adjustments, which severely restricts its ability to respond to country-specific variables, like in this case domestic inflation. Under these circumstances, a coordination of monetary policies to implement the fixed exchange rate raises overall macroeconomic stability and therefore welfare. Monetary unification provides such a coordination device (Cooley and Quadrini 2003; Pappa 2004).

Note that in this model the benefit of monetary unification compared with a unilateral fixed exchange rate regime does not derive from a credibility gain. By abstracting from speculative attacks, it is implicitly assumed that the fixed exchange rate is perfectly credible under both regimes. In reality, of course, a monetary union provides a much more credible fixed exchange rate regime than an exchange rate peg, due to the much greater costs of leaving or dissolving a monetary union (see, e.g., Eichengreen 1993). This credibility gain adds to the coordination gain described above.

Robustness. Under local-currency pricing or under cost-push shocks, the results are qualitatively very similar (see appendix C, figures C.4–C.6). The MU regime tends to be superior to unilateral exchange rate targeting but inferior to bilateral exchange rate targeting. Again, local-currency pricing and cost-push shocks work in favor of the MU regime, for the reasons explained above. Notably, under cost-push shocks, the MU regime and the bilateral exchange rate targeting regime are nearly identical for most values of the exchange rate coefficient.

3.4 Interest Rate Smoothing

Finally, in this subsection, monetary policy engages in interest rate smoothing. Under the FX regime and PPI inflation targeting, the interest rate rules for both countries are given by

$$\hat{R}_t = \phi_R \hat{R}_{t-1} + (1 - \phi_R)\phi_\pi \pi_{H,t} \tag{67}$$

$$\hat{R}^*_t = \phi_R \hat{R}^*_{t-1} + (1 - \phi_R)\phi_\pi \pi^*_{F,t}. \tag{68}$$

Under CPI inflation targeting, they take the following form:

$$\hat{R}_t = \phi_R \hat{R}_{t-1} + (1 - \phi_R)\phi_\pi \pi_t \tag{69}$$

$$\hat{R}^*_t = \phi_R \hat{R}^*_{t-1} + (1 - \phi_R)\phi_\pi \pi^*_t. \tag{70}$$

Under the MU regime, the interest rate rule of the common monetary policy is given by:

$$\hat{R}^{MU}_t = \phi_R \hat{R}^{MU}_{t-1} + (1 - \phi_R)\phi_\pi \pi^{MU}_t. \tag{71}$$

In all of these cases, the inflation coefficient ϕ_π is set to 1.5.

Producer-Currency Pricing and Productivity Shocks. Whether countries are better off under the MU regime or under the FX regime depends on the degree of interest rate smoothing implemented by monetary policy, which is particularly true under PPI inflation targeting (figure 6, solid blue and dashed red line).[20] Starting with very low degrees of interest rate smoothing (i.e., low values of ϕ_R), the two countries are better off under the MU regime. As the degree of interest rate smoothing increases, the welfare loss decreases faster under the FX regime than under the MU regime. At some point, the welfare ranking changes and the two countries are better off under the FX regime.

As described in section 3.1, the MU regime entails the cost of higher instability of both the output gap and the terms-of-trade gap, but the benefit of higher stability of the PPI inflation rate. This is again due to the mechanism mentioned earlier: As the nominal exchange rate is fixed and prices are sticky, the terms of trade

[20] For color versions of the figures, see the paper on the IJCB website (http://www.ijcb.org).

Figure 6. Welfare Loss as a Function of the Interest Rate Smoothing Coefficient (ϕ_R) under Producer-Currency Pricing and Productivity Shocks

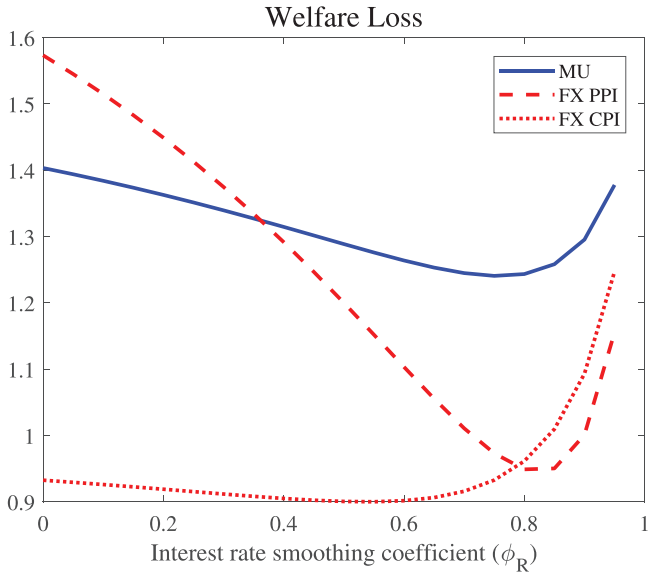


exhibit an inertial or history-dependent behavior, even if monetary policy does not smooth interest rates. This history dependence has the advantage of stabilizing private-sector expectations about future inflation and thereby actual inflation.

If monetary policy does not smooth interest rates under the FX regime, there is no such history dependence. The regime suffers from a kind of stabilization bias. As a result, PPI inflation is less stable under the FX regime. However, if monetary policy starts to smooth interest rates, it induces history dependence into the economy, with the same advantageous effect on inflation expectations. This effect strengthens as the degree of interest rate smoothing increases. If the degree of interest rate smoothing is sufficiently high, PPI inflation is more stable under the FX regime.

Under CPI inflation targeting, the degree of interest rate smoothing does not have such an important effect on the welfare ranking between the MU and FX regime (figure 6, solid blue and dotted red line). This is because, like the MU regime, the FX regime under CPI

inflation targeting features history dependence even if monetary policy does not smooth interest rates. As a result, engaging in interest rate smoothing, thereby inducing greater history dependence into the economy, does not change the relative welfare performance of the FX and MU regimes dramatically.[21]

Robustness. The results continue to hold under the combination of local-currency pricing and productivity shocks (see appendix C, figure C.7). As before, the differences in welfare losses between the MU and the FX regime are smaller because exchange rate flexibility is less beneficial under local-currency pricing due to the missing effect on international relative goods prices.

Under cost-push shocks, while the welfare performance of the MU regime relative to the FX regime continues to deteriorate with the degree of interest rate smoothing under PPI inflation targeting, there is no longer a change in the ranking, at least under the baseline calibration. For very high degrees of interest rate smoothing, the welfare loss is basically identical under both currency regimes. Under CPI inflation targeting, the degree of interest rate smoothing continues to have a much more limited influence on the welfare implications of monetary unification, as was the case under productivity shocks. But since cost-push shocks work in its favor, the likelihood of monetary unification being beneficial is higher than under productivity shocks.

4. Monetary Policy and Trade Openness

The conduct of monetary policy is not only an independent OCA criterion by itself, as illustrated in section 3, but it can also modify the nature of other OCA criteria. This is demonstrated in the following using the degree of trade openness as an example. But first, I briefly summarize how the relationship between trade openness and the costs and benefits of a monetary union is seen in the literature.

[21] As evident from figure 6, the welfare ranking does not change at all with the degree of interest rate smoothing under CPI inflation targeting. For other parameter constellations, however, this is the case, e.g., if $\phi_\pi = 1.2$ instead of 1.5.

4.1 Trade Openness in OCA Theory

The degree of trade openness or trade integration is one of the oldest and most prominent OCA criteria. Most studies have established a positive link between trade openness and the likelihood of a monetary union being beneficial. More precisely, the more open economies are, the smaller are the costs and the larger are the benefits associated with monetary unification.

McKinnon (1963) first proposed trade openness as an OCA criterion. He argued that with an increasingly open economy, the effects of exchange rate fluctuations on consumer prices via import prices become greater, thereby making it more difficult for monetary policy to maintain (consumer) price stability. Thus, the costs of giving up monetary independence decrease with the degree of trade openness.

One of the main costs attributed to monetary unification is the loss of the ability to react to asymmetric (i.e., country-specific) shocks via monetary policy and the nominal exchange rate. However, there are conflicting views on whether the incidence of country-specific shocks decreases or increases with the degree of trade openness. This depends on whether trade between countries is characterized predominantly by intra-industry trade or inter-industry trade. In the former case, industry-specific shocks affect countries symmetrically, thus an increase in the degree of trade openness reduces the cost of giving up monetary independence (Emerson et al. 1992, chapter 6.2). In the latter case, industry-specific shocks affect countries asymmetrically, thus an increase in the degree of trade openness raises the cost of giving up monetary independence (Krugman 1991, p. 82).[22]

The benefits traditionally associated with monetary unification are usually considered to increase with the degree of trade openness, such as the elimination of transaction costs when exchanging currencies, the increase in price transparency across countries, or the elimination of exchange rate risk (e.g., De Grauwe 2012, chapter 3.8). The latter point is also made by Kollmann (2004) using a New Keynesian DSGE model similar to that employed in the present study. He concludes that if the nominal exchange rate is subject to shocks,

[22]See De Grauwe (2012, chapter 2.1) for a more detailed description and assessment of the "European Commission view" and the "Krugman view."

a monetary union is welfare improving because the cost of giving up monetary independence is overcompensated by the benefit of eliminating exchange rate risk. This benefit increases with the degree of trade openness because exchange rate shocks are more harmful to macroeconomic stability as economies become more open.

Furthermore, the coordination gains associated with monetary unification also increase with the degree of trade openness, as shown by Pappa (2004) using a similar model. Compared with a flexible exchange rate regime where the monetary authorities do not cooperate to maximize welfare, forming a monetary union eliminates the possibility of strategic terms-of-trade manipulations. This benefit increases with the degree of trade openness because terms-of-trade movements have larger effects on macroeconomic stability as economies become more open.

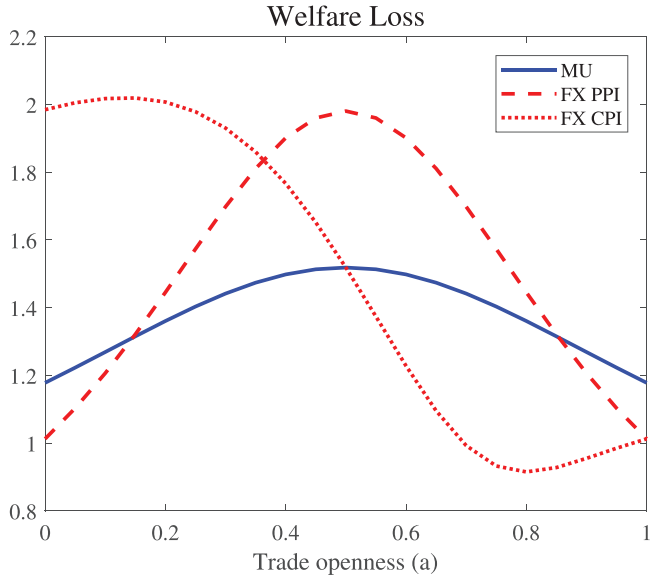
4.2 Monetary Policy and the Nature of Trade Openness as an OCA Criterion

The preceding overview shows that OCA theory mainly establishes a favorable relationship between the degree of trade openness and the costs and benefits of a monetary union. As shown next, this is highly sensitive to the way monetary policy is conducted.

In what follows, the interest rate rules are given by equations (46) through (50), with the inflation coefficient ϕ_π set to 1.5 in all cases.

Producer-Currency Pricing and Productivity Shocks. The influence of monetary policy on the nature of the degree of trade openness as an OCA criterion is particularly clear when distinguishing between PPI and CPI inflation targeting. First consider the case of PPI inflation targeting (figure 7, solid blue and dashed red line). Two observations are noteworthy. First, under both the MU and the FX regime, the relationship between the welfare loss and the degree of trade openness is symmetric around a trade-to-GDP ratio of 100 percent ($a = 1/2$). Second, the two countries are better off under the FX regime if they are either relatively closed (a close to one) or very open to trade (a close to zero), but better off under the MU regime for intermediate values. Thus, the likelihood of the MU regime being beneficial first increases and then decreases with the degree of trade openness.

Figure 7. Welfare Loss as a Function of the Degree of trade Openness (a) under Producer-Currency Pricing and Productivity Shocks

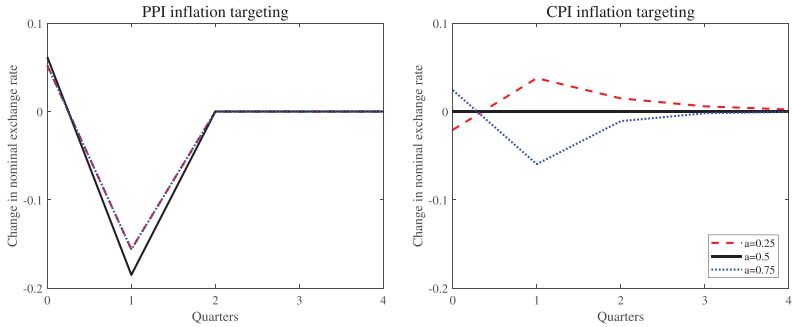


The relationship between trade openness and the welfare ranking between the MU and FX regime changes considerably if monetary policy targets CPI inflation rates instead of PPI inflation rates (figure 7, solid blue and dotted red line). First, the relationship between the welfare loss and the degree of trade openness is no longer symmetric under the FX regime.[23] Second, the two countries are better off under the FX regime for trade-to-GDP ratios between 0 and 100 percent ($1/2 < a \leq 1$) and better off under the MU regime for ratios between 100 and 200 percent ($0 \leq a < 1/2$). Thus, the likelihood of the MU regime being beneficial increases with the degree of trade openness.

The key to understanding these results is again the role played by the nominal exchange rate in stabilizing the terms-of-trade gap. Consider a positive productivity shock in country H. Recall that

[23]Recall that there is no difference between PPI and CPI inflation targeting under the MU regime.

Figure 8. Impulse Response of the Change in the Nominal Exchange Rate ($\Delta\hat{S}_t$) to a Positive One-Off Productivity Shock in Country H for Three Different Degrees of Trade Openness (a), with $\alpha = 0.2$, under Producer-Currency Pricing



Notes: Left panel: PPI inflation targeting. Right panel: CPI inflation targeting.

the efficient terms of trade unambiguously increase on impact (see equation (57)), thus an increase in the nominal exchange rate would help to stabilize the terms-of-trade gap, thereby reducing the welfare loss. But whether the nominal exchange rate stabilizes or destabilizes the terms-of-trade gap depends crucially on whether monetary policy targets PPI or CPI inflation.

Under PPI inflation targeting, the impact response of the nominal exchange rate is positive irrespective of the degree of trade openness, i.e., country H's currency depreciates (figure 8, left panel).[24] Thus, the nominal exchange rate pushes the sticky-price terms of trade in the same direction as the efficient terms of trade, thereby stabilizing the terms-of-trade gap to some extent. Note that the response of the nominal exchange rate is identical for $\alpha = 0.25$ and $\alpha = 0.75$, which explains the symmetric pattern visible in figure 7.

By contrast, under CPI inflation targeting, the impact response of the nominal exchange rate is positive if the two countries have a trade-to-GDP ratio below 100 percent ($a > 1/2$), but negative if

[24]The degree of price stickiness was set to a low value ($\alpha = 0.2$) to ensure that the differences in the impulse responses are clearly visible. The differences for higher degrees of price stickiness are smaller but qualitatively the same.

it is above 100 percent ($a < 1/2$) (figure 8, right panel). Thus, the nominal exchange rate helps to stabilize the terms-of-trade gap only in the first case. In the latter case, the nominal exchange rate actually destabilizes the terms-of-trade gap by pushing the sticky-price terms of trade away from the efficient terms of trade.

This is robust with respect to the other deep parameters of the economy.[25] To see this, insert the country-specific interest rate rules (48) and (49) together with the definitions of the CPI inflation rates (10) and (11) and the terms-of-trade identity (9) into the uncovered interest parity condition (3) to obtain

$$\Delta\hat{S}_t = (2a-1)\Delta\hat{T}_t + \frac{1}{\phi_\pi}E_t\Delta\hat{S}_{t+1}. \tag{72}$$

Solving forward yields

$$\Delta\hat{S}_t = (2a-1)E_t\sum_{k=0}^{\infty}\left(\frac{1}{\phi_\pi}\right)^k\Delta\hat{T}_{t+k}. \tag{73}$$

Accordingly, the current change in the nominal exchange rate depends on the discounted sum of current and expected future changes in the terms of trade. Importantly, this relationship is positive if $a > 1/2$, but negative if $a < 1/2$.

Under PPI inflation targeting, the analogous equations are given by

$$\Delta\hat{S}_t = \Delta\hat{T}_t + \frac{1}{\phi_\pi}E_t\Delta\hat{S}_{t+1} \tag{74}$$

and

$$\Delta\hat{S}_t = E_t\sum_{k=0}^{\infty}\left(\frac{1}{\phi_\pi}\right)^k\Delta\hat{T}_{t+k}. \tag{75}$$

[25] In particular, it does not make a difference whether $\rho\theta$ is smaller than, equal to, or larger than 1, although this condition has important macroeconomic implications. For example, it determines whether the cross-country correlation of output is positive, zero, or negative (see, e.g., Corsetti, Dedola, and Leduc 2011 for details). Also, if it is zero ($\rho\theta = 1$), the terms-of-trade gap vanishes from the welfare loss function (41).

In contrast to the CPI inflation targeting case, the relationship between the current change in the nominal exchange rate and the discounted sum of current and expected future changes in the terms of trade is always positive, regardless of the degree of trade openness a.[26]

The intuition behind the fact that the nominal exchange rate can be destabilizing under CPI inflation targeting is the following. If the trade-to-GDP ratio is above 100 percent ($a < 1/2$), consumer prices in one country are determined mainly by producer prices in the other country because consumers consume more imported goods than home-produced goods. If monetary policy targets consumer prices, interest rate adjustments in one country are triggered mainly by producer price changes in the other country. This pushes the nominal exchange rate, which depends on the interest rate differential between the two countries, away from the efficient terms of trade. As a result, the welfare-relevant terms-of-trade gap is destabilized by the nominal exchange rate. Under these circumstances, a fixed exchange rate would make the countries better off because this is neither destabilizing nor stabilizing. For this reason, the countries are better off under the MU regime for $a < 1/2$.

In the special case of a trade-to-GDP ratio of exactly 100 percent ($a = 1/2$), the two countries are indifferent between the FX and the MU regime under CPI inflation targeting. This is because the nominal exchange rate is constant under both regimes.[27] Under the MU regime, the nominal exchange rate is fixed by construction. Under the FX regime, it is fixed by coincidence. That is, by targeting CPI inflation rates, the two countries unintentionally implement a symmetric fixed exchange rate regime. This is because consumer price changes and thus interest rate adjustments are identical in the two countries.

Lastly, as shown, under PPI inflation targeting the nominal exchange rate stabilizes the terms-of-trade gap regardless of the degree of trade openness. Nonetheless, for a broad range of degrees of

[26] Note that equation (73) and equation (75) are equivalent if $a = 1$. In this case, there is no difference between PPI and CPI inflation targeting. This is because the consumer price index equals the producer price index if $a = 1$; see equations (10) and (11).

[27] According to equation (73), $\Delta\hat{S}_t = 0$ if $a = 1/2$.

trade openness, the FX regime is inferior to the MU regime, where the nominal exchange rate is fixed (recall figure 7). This is again due to the inherent benefit of monetary unions explained earlier: As the nominal exchange rate is fixed and prices are sticky, the terms of trade exhibit an inertial or history-dependent behavior. This history dependence has the advantage of stabilizing private-sector expectations about future inflation and thereby actual inflation. This benefit weakens as the degree of trade openness becomes either very low ($a \to 1$) or very high ($a \to 0$). In the extreme cases, consumers consume only one of the two internationally traded goods. The relative price (terms of trade) becomes irrelevant for price setters, and the terms of trade no longer affect inflation.[28] As a result, the inertia in the terms of trade no longer has a stabilizing effect on inflation.

Robustness. The conclusion that the nature of trade openness as an OCA criterion differs markedly between PPI and CPI inflation targeting is robust to local-currency pricing or cost-push shocks, though how that difference specifically looks varies from case to case (see appendix C, figure C.8). The only exception to this conclusion results if local-currency pricing and productivity shocks concur. In that case, there is no difference between PPI and CPI inflation targeting in the sense that the likelihood of the MU regime being beneficial is lowest under either very closed or very open economies.

5. Conclusion

The costs and benefits of moving from a flexible exchange rate regime to a monetary union depend critically on the conduct of monetary policy. Whether countries are better off in one or the other currency regime is sensitive not only to the choice of the variables that monetary policy targets but also to the strength of the response to these target variables. In particular, when monetary policy in each country responds to inflation aggressively or implements a high degree of interest rate smoothing, forming a monetary union, where the common monetary authority continues to follow the same policy, tends

[28]Note how the terms of trade vanish from the New Keynesian Phillips curves (7) and (8) if $a = 0$ or $a = 1$.

to make countries worse off in terms of welfare by reducing macroeconomic stability. By contrast, when monetary policy responds to inflation only modestly or implements a low degree of interest rate smoothing, forming a monetary union tends to make countries better off. Furthermore, monetary unification is beneficial when monetary policy responds to *output*, whereas it is costly when monetary policy responds to the *output gap*. And finally, it is important whether countries respond to the nominal exchange rate and whether they do so in a coordinated or uncoordinated way. In the latter case, monetary unification is generally beneficial, whereas it is costly in the former case.

In addition to being an OCA criterion itself, monetary policy has the potential to modify the nature of traditional OCA criteria, such as the degree of trade openness. Whether the likelihood of a monetary union being beneficial *increases* with the degree of trade openness, as proposed by the vast bulk of OCA studies, depends critically on whether monetary policy targets producer price inflation or consumer price inflation. In the former case, it is possible that the likelihood of a monetary union being beneficial *decreases* with the degree of trade openness.

With few exceptions, these conclusions are not to any important degree sensitive to the price-setting assumption (producer-currency pricing versus local-currency pricing) or the type of shocks (productivity shocks versus cost-push shocks). However, local-currency pricing and cost-push shocks—individually as well as jointly—tend to increase the likelihood that countries benefit from monetary unification.

Appendix A. Flexible Exchange Rate Regime

This appendix contains the full derivation of the model under the flexible exchange rate regime for producer-currency pricing and local-currency pricing, respectively (based on Corsetti, Dedola, and Leduc 2011). The world, which consists of two countries labeled H and F, is populated by a continuum of agents on the interval $[0, 1]$. The population on the segment $[0, n)$ lives in country H; the population on the segment $[n, 1]$ lives in country F. Thus, n measures the population size as a fraction of world population. An agent is both

consumer and producer. He produces a single differentiated good and consumes all the goods produced in both countries.

A.1 Consumer Problem

Agent j in country H derives positive utility from consumption C^j and negative utility from producing the differentiated good $y(h)$. The present discounted value of lifetime utility U^j is given by[29]

$$U^j = E_0 \sum_{t=0}^{\infty} \beta^t \left[\zeta_{C,t} \frac{{C_t^j}^{1-\rho} - 1}{1-\rho} - \zeta_{Y,t}^{-\eta} \frac{y_t(h)^{1+\eta}}{1+\eta} \right]. \tag{A.1}$$

E denotes the expectations operator, β the discount factor, ρ the inverse of the intertemporal elasticity of substitution in consumption, and η the inverse of the elasticity of producing the differentiated good.[30] $\zeta_{Y,t}$ and $\zeta_{C,t}$ denote shocks to productivity and to preferences in consumption, respectively. These shocks are common to all agents living in country H.

Consumption Preferences. The agent consumes a bundle of differentiated goods both from country H and from country F according to the following constant-elasticity-of-substitution (CES) aggregator:

$$C_t^j = \left[a^{\frac{1}{\theta}} {C_{H,t}^j}^{\frac{\theta-1}{\theta}} + (1-a)^{\frac{1}{\theta}} {C_{F,t}^j}^{\frac{\theta-1}{\theta}} \right]^{\frac{\theta}{\theta-1}}, \tag{A.2}$$

where the bundles of differentiated goods are given by aggregators according to Dixit and Stiglitz (1977):

$$C_{H,t}^j = \left[\left(\frac{1}{n}\right)^{\frac{1}{\sigma}} \int_0^n c_t^j(h)^{\frac{\sigma-1}{\sigma}} dh \right]^{\frac{\sigma}{\sigma-1}} \tag{A.3}$$

$$C_{F,t}^j = \left[\left(\frac{1}{1-n}\right)^{\frac{1}{\sigma}} \int_n^1 c_t^j(f)^{\frac{\sigma-1}{\sigma}} df \right]^{\frac{\sigma}{\sigma-1}}.$$

[29] In Corsetti, Dedola, and Leduc (2011), the agent derives utility also from the liquidity services of holding money. I abstract from money in the utility function, since monetary policy is conducted via interest rate rules.

[30] The parameter η is equivalent to the inverse of the Frisch elasticity of labor supply.

These preferences imply (i) that the elasticity of substitution between differentiated goods c_t^j from one country is σ, which is assumed to be greater than one and equal across countries, (ii) that the elasticity of substitution between the bundles of goods from the two countries $C_{H,t}$ and $C_{F,t}$ is θ, which is assumed to be greater than zero and equal across countries, and (iii) that the steady-state share of imported goods in overall consumption expenditures is $1-a$. If $a > 1/2$, the agent consumes more goods from the country the agent lives in than from the other country, i.e., the agent has a home bias in consumption. This home bias is assumed to be symmetric across countries. Thus, the CES aggregator for an agent j living in country F is given by

$$C_t^{j*} = \left[(1-a)^{\frac{1}{\theta}} C_{H,t}^{j*}{}^{\frac{\theta-1}{\theta}} + a^{\frac{1}{\theta}} C_{F,t}^{j*}{}^{\frac{\theta-1}{\theta}}\right]^{\frac{\theta}{\theta-1}}. \tag{A.4}$$

Accordingly, the consumer price index (CPI) in country H expressed in country H's currency is given by

$$P_t = \left[a P_{H,t}{}^{1-\theta} + (1-a) P_{F,t}{}^{1-\theta}\right]^{\frac{1}{1-\theta}}, \tag{A.5}$$

where the producer price indexes (PPI) for the bundles of differentiated goods expressed in country H's currency are defined by

$$P_{H,t} = \left[\frac{1}{n}\int_0^n p_t(h)^{1-\sigma} dh\right]^{\frac{1}{1-\sigma}} \tag{A.6}$$

$$P_{F,t} = \left[\frac{1}{1-n}\int_n^1 p_t(f)^{1-\sigma} df\right]^{\frac{1}{1-\sigma}}.$$

The CPI in country F expressed in country F's currency is given by

$$P_t^* = \left[(1-a) P_{H,t}^*{}^{1-\theta} + a P_{F,t}^*{}^{1-\theta}\right]^{\frac{1}{1-\theta}}. \tag{A.7}$$

Producer-Currency Pricing. In their role as producers, agents charge one price for their good irrespective of whether the good is sold in their country or is exported to the other country, setting the price in their country's currency. Furthermore, exporting

does not entail transportation costs. These assumptions imply that the law of one price holds, i.e., a single differentiated good has the same price in both countries if expressed in the same currency, and that exchange rate pass-through is complete:

$$p_t(h) = S_t p_t^*(h), \qquad p_t(f) = S_t p_t^*(f), \tag{A.8}$$

where $p_t(h)$ denotes the price of a differentiated good $y(h)$ produced in country H denominated in country H's currency, $p_t^*(h)$ denotes the price of the same good $y(h)$ denominated in country F's currency, $p_t(f)$ denotes the price of a differentiated good $y(f)$ produced in country F denominated in country H's currency, $p_t^*(f)$ denotes the price of the same good $y(f)$ denominated in country F's currency, and S_t is the nominal exchange rate defined as the price of country F's currency in terms of country H's currency. Given equations (A.6), it is straightforward to show that the law of one price for each differentiated good translates into the law of one price for each bundle of goods:

$$P_{H,t} = S_t P_{H,t}^*, \qquad P_{F,t} = S_t P_{F,t}^*. \tag{A.9}$$

In general, the law of one price does not translate into purchasing power parity. Thus, the real exchange rate, defined as the ratio of country-specific consumer prices

$$Q_t = \frac{S_t P_t^*}{P_t}, \tag{A.10}$$

adjusts in response to changing economic conditions. Purchasing power parity ($Q_t = 1$) only holds if the consumption baskets are identical across countries ($a = 1/2$).

Another international relative price of interest are the terms of trade, defined from the perspective of country H as the ratio of the price of imported goods to the price of exported goods:

$$T_t = \frac{P_{F,t}}{S_t P_{H,t}^*}. \tag{A.11}$$

Under producer-currency pricing, where the law of one price holds, the terms of trade can be expressed as

$$T_t = \frac{S_t P_{F,t}^*}{P_{H,t}}. \tag{A.12}$$

Agent j in country H takes three decisions with respect to his consumption choices. First, he decides on the overall level of consumption C_t^j. Second, given C_t^j, the agent optimally allocates expenditures between the bundles of differentiated goods $C_{H,t}^j$ and $C_{F,t}^j$ by minimizing total expenditure $P_t C_t^j$ subject to the CES aggregator (A.2). As a result, demand for these bundles is given by

$$C_{H,t}^j = a\left(\frac{P_{H,t}}{P_t}\right)^{-\theta} C_t^j, \qquad C_{F,t}^j = (1-a)\left(\frac{P_{F,t}}{P_t}\right)^{-\theta} C_t^j. \tag{A.13}$$

Third, given $C_{H,t}^j$ and $C_{F,t}^j$, the agent optimally allocates expenditures between the differentiated goods by minimizing $P_{H,t}C_{H,t}^j$ and $P_{F,t}C_{F,t}^j$ subject to equations (A.3). This yields

$$c_t^j(h) = \frac{1}{n}\left(\frac{p_t(h)}{P_{H,t}}\right)^{-\sigma} C_{H,t}^j, \qquad c_t^j(f) = \frac{1}{1-n}\left(\frac{p_t(f)}{P_{F,t}}\right)^{-\sigma} C_{F,t}^j. \tag{A.14}$$

Combining (A.13) and (A.14) yields

$$\begin{aligned} c_t^j(h) &= \frac{a}{n}\left(\frac{p_t(h)}{P_{H,t}}\right)^{-\sigma}\left(\frac{P_{H,t}}{P_t}\right)^{-\theta} C_t^j \\ c_t^j(f) &= \frac{1-a}{1-n}\left(\frac{p_t(f)}{P_{F,t}}\right)^{-\sigma}\left(\frac{P_{F,t}}{P_t}\right)^{-\theta} C_t^j. \end{aligned} \tag{A.15}$$

Analogously, demand equations for an agent j in country F are given by

$$\begin{aligned} c_t^{j*}(h) &= \frac{1-a}{n}\left(\frac{p_t^*(h)}{P_{H,t}^*}\right)^{-\sigma}\left(\frac{P_{H,t}^*}{P_t^*}\right)^{-\theta} C_t^{j*} \\ c_t^{j*}(f) &= \frac{a}{1-n}\left(\frac{p_t^*(f)}{P_{F,t}^*}\right)^{-\sigma}\left(\frac{P_{F,t}^*}{P_t^*}\right)^{-\theta} C_t^{j*}. \end{aligned} \tag{A.16}$$

Market clearing for the differentiated goods $y_t(h)$ and $y_t(f)$ requires

$$y_t(h) = \int_0^n c_t^j(h)dj + \int_n^1 c_t^{j*}(h)dj \tag{A.17}$$
$$y_t(f) = \int_0^n c_t^j(f)dj + \int_n^1 c_t^{j*}(f)dj.$$

Using equations (A.15) and (A.16), world demand for the differentiated goods can be expressed as

$$y_t(h) = \frac{1}{n}\left(\frac{p_t(h)}{P_{H,t}}\right)^{-\sigma}\left(\frac{P_{H,t}}{P_t}\right)^{-\theta}\left[aC_t + (1-a)Q_t^\theta C_t^*\right] \tag{A.18}$$
$$y_t(f) = \frac{1}{1-n}\left(\frac{p_t(f)}{P_{F,t}}\right)^{-\sigma}\left(\frac{P_{F,t}}{P_t}\right)^{-\theta}\left[(1-a)C_t + aQ_t^\theta C_t^*\right],$$

where $C_t \equiv \int_0^n C_t^j dj$ and $C_t^* \equiv \int_n^1 C_t^{j*} dj$.

Similar to (A.3), aggregate output in the two countries can be expressed as

$$Y_{H,t} = \left[\left(\frac{1}{n}\right)^{\frac{1}{\sigma}}\int_0^n y_t(h)^{\frac{\sigma-1}{\sigma}}dh\right]^{\frac{\sigma}{\sigma-1}} \tag{A.19}$$
$$Y_{F,t} = \left[\left(\frac{1}{1-n}\right)^{\frac{1}{\sigma}}\int_n^1 y_t(f)^{\frac{\sigma-1}{\sigma}}df\right]^{\frac{\sigma}{\sigma-1}}.$$

Inserting equations (A.18) into the previous equations finally yields aggregate demand

$$Y_{H,t} = \left(\frac{P_{H,t}}{P_t}\right)^{-\theta}\left[aC_t + (1-a)Q_t^\theta C_t^*\right] \tag{A.20}$$
$$Y_{F,t} = \left(\frac{P_{F,t}}{P_t}\right)^{-\theta}\left[(1-a)C_t + aQ_t^\theta C_t^*\right].$$

Local-Currency Pricing. Under local-currency pricing, agents set the price for their good in the currency of the destination market, i.e., in their country's currency if the good is sold domestically,

and in the other country's currency if the good is sold abroad. When able to change prices in a given period, the agent will reset in both currencies. Since this is not possible every period, subsequent fluctuations in the nominal exchange rate lead to temporary deviations from the law of one price.

Following the same steps as before, aggregate demand under local-currency pricing is given by

$$Y_{H,t} = \left(\frac{P_{H,t}}{P_t}\right)^{-\theta} \left[aC_t + (1-a)\left(\frac{Q_t}{\Delta_{H,t}}\right)^{\theta} C_t^*\right] \tag{A.21}$$

$$Y_{F,t} = \left(\frac{P_{F,t}}{P_t}\right)^{-\theta} \left[(1-a)C_t + a\left(\frac{Q_t}{\Delta_{F,t}}\right)^{\theta} C_t^*\right],$$

where

$$\Delta_{H,t} = \frac{S_t P_{H,t}^*}{P_{H,t}}, \qquad \Delta_{F,t} = \frac{S_t P_{F,t}^*}{P_{F,t}} \tag{A.22}$$

represent the deviation from the law of one price for each good.

Cross-Country Risk Sharing. Asset markets are assumed to be complete within and across countries. Agents can insure against all possible states of nature by holding a portfolio of state-contingent, one-period securities whose real value (denominated in units of the consumption-based price index) is denoted by $B_t^{H,j}$ and whose vector of prices is denoted by q_t^H. In addition, agents can trade in a noncontingent, one-period bond whose nominal value (denominated in country H's currency) is denoted by B_t^j and whose nominal interest rate is denoted by R_t. The intertemporal budget constraint of agent j in country H is then given by

$$C_t^j + q_t^H B_t^{H,j} + \frac{B_t^j}{P_t(1+R_t)} = B_{t-1}^{H,j} + \frac{B_{t-1}^j}{P_t} + (1-\tau_t^H)\frac{p_t(h)y_t(h)}{P_t}. \tag{A.23}$$

The agent's income stems also from sales revenues $p_t(h)y_t(h)$ net of a proportional, country-specific tax τ_t^H.[31]

[31] The tax will turn out to be a subsidy to exactly offset the distortion caused by monopolistic competition.

All contingent securities and noncontingent bonds are assumed to be in zero supply in the initial period, so $B_0^{H,j} = B_0^j = 0$ for all j. Together with the facts that, within countries, agents have identical preferences and that asset markets are complete, this implies perfect risk sharing of consumption within each country. Therefore, it is possible to analyze the consumer problem from the viewpoint of the representative agent of country H and country F.

The representative agent in country H maximizes his lifetime utility (A.1) subject to the budget constraint (A.23). By combining the resulting first-order conditions with respect to consumption and bond holdings, the usual Euler consumption equation is then given by

$$U_C(C_t, \zeta_{C,t}) = (1 + R_t)\beta E_t \left\{ U_C(C_{t+1}, \zeta_{C,t+1}) \frac{P_t}{P_{t+1}} \right\}. \tag{A.24}$$

The Euler consumption equation for the representative agent in country F is obtained analogously and given by

$$U_C(C_t^*, \zeta_{C,t}^*) = (1 + R_t^*)\beta E_t \left\{ U_C(C_{t+1}^*, \zeta_{C,t+1}^*) \frac{P_t^*}{P_{t+1}^*} \right\}. \tag{A.25}$$

Complete asset markets across countries leads to price equalization in the state-contingent securities (expressed in country H's currency), implying the following risk-sharing condition:

$$\beta \frac{U_C(C_{t+1}, \zeta_{C,t+1})}{U_C(C_t, \zeta_{C,t})} \frac{P_t}{P_{t+1}} = \beta \frac{U_C(C_{t+1}^*, \zeta_{C,t+1}^*)}{U_C(C_t^*, \zeta_{C,t}^*)} \frac{S_t P_t^*}{S_{t+1} P_{t+1}^*}. \tag{A.26}$$

Inserting the country-specific Euler consumption equations yields the uncovered interest parity condition, according to which the expected change in the nominal exchange rate corresponds to the ratio of the country-specific interest rates:

$$E_t \Delta S_{t+1} = \frac{1 + R_t}{1 + R_t^*}. \tag{A.27}$$

Assuming net foreign asset positions to be initially symmetric and applying the definition of the real exchange rate (A.10), the risk-sharing condition takes the following form:

$$Q_t = \left(\frac{C_t^*}{C_t} \right)^{-\rho} \frac{\zeta_{C,t}^*}{\zeta_{C,t}}. \tag{A.28}$$

A.2 Producer Problem

In their role as producers, agents act in an environment of monopolistic competition, in which they dispose of some degree of market power. Furthermore, prices are sticky in the sense that the agent is able to change the price in a given period with a fixed probability, as in (A.10). The probability of being able to change the price is identical across countries and given by $1-\alpha$.

Producer-Currency Pricing. Agent j in country H maximizes expected, discounted profits by choosing the price $\tilde{p}_t(h)$ taking into account that demand depends on the chosen price and that the price may remain unchanged for some periods in the domestic market. In the foreign market, the price in foreign currency fluctuates with the nominal exchange rate, so that the law of one price always holds (complete exchange rate pass-through). Formally, the agent maximizes

$$E_t \sum_{k=0}^{\infty} (\alpha\beta)^k \left[\lambda_{t+k}(1-\tau_t^H)\tilde{p}_t(h)\tilde{y}_{t,t+k}(h) - V(\tilde{y}_{t,t+k}(h), \zeta_{Y,t+k})\right] \tag{A.29}$$

subject to the demand function

$$\tilde{y}_{t,t+k}(h) = \frac{1}{n}\left(\frac{\tilde{p}_t(h)}{P_{H,t+k}}\right)^{-\sigma}\left(\frac{P_{H,t+k}}{P_{t+k}}\right)^{-\theta} \times \left[aC_{t+k} + (1-a)Q_{t+k}^{\theta}C_{t+k}^{*}\right], \tag{A.30}$$

where $\tilde{y}_{t,t+k}(h)$ denotes total demand of good h at time $t+k$ if the price $\tilde{p}_t(h)$ prevails. Profits are expressed in utility units. Therefore, nominal sales revenues net of taxes $(1-\tau_{t+k}^H)\tilde{p}_t(h)\tilde{y}_{t,t+k}(h)$ are converted into utility units using the marginal utility of nominal revenues $\lambda_{t+k} = \frac{U_C(C_{t+k},\zeta_{C,t+k})}{P_{t+k}}$. The cost of production expressed in utility units is given by the function $V = \zeta_{Y,t}^{-\eta}\frac{y_t(h)^{1+\eta}}{1+\eta}$.

The first-order condition yields the optimal price

$$\tilde{p}_t(h) = \frac{E_t \sum_{k=0}^{\infty}(\alpha\beta)^k \frac{\sigma}{(\sigma-1)(1-\tau_{t+k}^H)} V_y(\tilde{y}_{t,t+k}(h), \zeta_{Y,t+k})\tilde{y}_{t,t+k}(h)}{E_t \sum_{k=0}^{\infty}(\alpha\beta)^k \lambda_{t+k}\tilde{y}_{t,t+k}(h)}, \tag{A.31}$$

where V_y denotes the derivative of function V with respect to output $\tilde{y}(h)$. All agents that live in the same country and are able to reset their price in a certain period will set the same price, since they share identical preferences (function V) and face the same demand curves, which depend only on aggregate variables such as P_H, P, P^*, S, C, and C^*, and the common elasticities of substitution σ and θ. Hence, in a given period, a fraction $1-\alpha$ of agents will set the same optimal price, while for a fraction α of agents the price from the previous period remains effective:

$$P_{H,t} = \left[\alpha P_{H,t-1}^{1-\sigma} + (1-\alpha)\tilde{p}_t(h)^{1-\sigma}\right]^{\frac{1}{1-\sigma}} \tag{A.32}$$

$$P_{F,t}^* = \left[\alpha P_{F,t-1}^{*1-\sigma} + (1-\alpha)\tilde{p}_t^*(f)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}.$$

When prices are flexible, the optimal-price equation (A.31) for country H simplifies to

$$\frac{P_{H,t}}{P_t} = \frac{\sigma}{(\sigma-1)(1-\tau_t^H)} \frac{V_y(y_t(h), \zeta_{Y,t})}{U_C(C_t, \zeta_{C,t})}, \tag{A.33}$$

and for country F to

$$\frac{P_{F,t}^*}{P_t^*} = \frac{\sigma}{(\sigma-1)(1-\tau_t^F)} \frac{V_y(y_t(f), \zeta_{Y,t}^*)}{U_C(C_t^*, \zeta_{C,t}^*)}. \tag{A.34}$$

The markup that agents in country $i = H, F$ are able to charge is defined as

$$\mu_t^i \equiv \frac{\sigma}{(\sigma-1)(1-\tau_t^i)}. \tag{A.35}$$

Local-Currency Pricing. In contrast to producer-currency pricing, agent j in country H sets two prices to maximize expected, discounted profits: the price in country H's currency for sales in country H and the price in country F's currency for sales in country F. Formally, the agent maximizes

$$E_t \sum_{k=0}^{\infty} (\alpha\beta)^k \left[\lambda_{t+k}(1-\tau_t^H)\left[\tilde{p}_t(h)\tilde{c}_{t,t+k}(h) + S_{t+k}\tilde{p}_t^*(h)\tilde{c}_{t,t+k}^*(h)\right]\right.$$
$$\left. -V(\tilde{y}_{t,t+k}(h), \zeta_{Y,t+k})\right] \tag{A.36}$$

subject to the demand functions

$$\tilde{c}_{t,t+k}(h) = \frac{1}{n}\left(\frac{\tilde{p}_t(h)}{P_{H,t+k}}\right)^{-\sigma}\left(\frac{P_{H,t+k}}{P_{t+k}}\right)^{-\theta} aC_{t+k} \tag{A.37}$$

$$\tilde{c}^*_{t,t+k}(h) = \frac{1}{n}\left(\frac{\tilde{p}^*_t(h)}{P^*_{H,t+k}}\right)^{-\sigma}\left(\frac{P^*_{H,t+k}}{P^*_{t+k}}\right)^{-\theta} (1-a)C^*_{t+k}.$$

The first-order conditions yield the two optimal prices

$$\tilde{p}_t(h) = \frac{E_t \sum_{k=0}^{\infty}(\alpha\beta)^k \mu^H_{t+k} V_y(\tilde{y}_{t,t+k}(h), \zeta_{Y,t+k})\tilde{c}_{t,t+k}(h)}{E_t \sum_{k=0}^{\infty}(\alpha\beta)^k \lambda_{t+k}\tilde{c}_{t,t+k}(h)} \tag{A.38}$$

$$\tilde{p}^*_t(h) = \frac{E_t \sum_{k=0}^{\infty}(\alpha\beta)^k \mu^H_{t+k} V_y(\tilde{y}_{t,t+k}(h), \zeta_{Y,t+k})\tilde{c}^*_{t,t+k}(h)}{E_t \sum_{k=0}^{\infty}(\alpha\beta)^k \lambda_{t+k} S_{t+k}\tilde{c}^*_{t,t+k}(h)}.$$

Together with the two optimal prices of agent j in country F, which are derived analogously, there are four sticky producer prices, compared with two in the case of producer-currency pricing. The corresponding producer price indexes are given by

$$P_{H,t} = \left[\alpha P^{1-\sigma}_{H,t-1} + (1-\alpha)\tilde{p}_t(h)^{1-\sigma}\right]^{\frac{1}{1-\sigma}} \tag{A.39}$$

$$P^*_{H,t} = \left[\alpha P^{*1-\sigma}_{H,t-1} + (1-\alpha)\tilde{p}^*_t(h)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}$$

$$P_{F,t} = \left[\alpha P^{1-\sigma}_{F,t-1} + (1-\alpha)\tilde{p}_t(f)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}$$

$$P^*_{F,t} = \left[\alpha P^{*1-\sigma}_{F,t-1} + (1-\alpha)\tilde{p}^*_t(f)^{1-\sigma}\right]^{\frac{1}{1-\sigma}}.$$

A.3 Log-Linearization

Deviations of the logarithm of a variable X_t from its steady state are denoted by $\hat{X}_t$ if prices are sticky and by $\tilde{X}^{fb}_t$ if prices are flexible and markups are neutralized (efficient allocation).

A.3.1 Sticky-Price Equilibrium under Producer-Currency Pricing

Under sticky prices and producer-currency pricing, the system of equations is given by

$$E_t\hat{C}_{t+1} = \hat{C}_t + \frac{1}{\rho}\left(\hat{R}_t - E_t\pi_{t+1} + E_t\hat{\zeta}_{C,t+1} - \hat{\zeta}_{C,t}\right) \tag{A.40}$$

$$\hat{Q}_t = \rho\left(\hat{C}_t - \hat{C}_t^*\right) + \left(\hat{\zeta}_{C,t}^* - \hat{\zeta}_{C,t}\right) \tag{A.41}$$

$$E_t\Delta\hat{S}_{t+1} = \hat{R}_t - \hat{R}_t^* \tag{A.42}$$

$$\hat{Q}_t = (2a-1)\hat{T}_t \tag{A.43}$$

$$\hat{Y}_{H,t} = 2a(1-a)\theta\hat{T}_t + a\hat{C}_t + (1-a)\hat{C}_t^* \tag{A.44}$$

$$\hat{Y}_{F,t} = -2a(1-a)\theta\hat{T}_t + (1-a)\hat{C}_t + a\hat{C}_t^* \tag{A.45}$$

$$\begin{aligned}\pi_{H,t} = {} & (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}_{H,t}^{fb}\right) - 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^H + \beta E_t\pi_{H,t+1}\end{aligned} \tag{A.46}$$

$$\begin{aligned}\pi_{F,t}^* = {} & (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}_{F,t}^{fb}\right) + 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^F + \beta E_t\pi_{F,t+1}^*\end{aligned} \tag{A.47}$$

$$\hat{T}_t = \hat{T}_{t-1} + \pi_{F,t}^* - \pi_{H,t} + \Delta\hat{S}_t \tag{A.48}$$

$$\pi_t = a\pi_{H,t} + (1-a)(\pi_{F,t}^* + \Delta\hat{S}_t) \tag{A.49}$$

$$\pi_t^* = (1-a)(\pi_{H,t} - \Delta\hat{S}_t) + a\pi_{F,t}^*. \tag{A.50}$$

Equation (A.40) is the log-linear approximation of the Euler consumption equation (A.24), where $\pi_t = \ln(P_t/P_{t-1})$. Equation (A.41) is the log-linear approximation of the risk-sharing condition (A.28), and equation (A.42) is the log-linear approximation of the uncovered interest parity condition (A.27).[32] Equation (A.43) describes

[32] Alternatively, the model could be specified by including both country-specific Euler consumption equations next to the risk-sharing condition, while omitting the uncovered interest parity condition.

the link between the real exchange rate and the terms of trade and is obtained by combining the log-linearized definition of the real exchange rate (A.10) with the log-linearized expressions for the country-specific CPIs (A.5) and (A.7), applying the law of one price. Equations (A.44) and (A.45) are obtained by log-linearizing the aggregate demand equations (A.20), using equation (A.43) to eliminate the real exchange rate.

Equations (A.46) and (A.47) represent the New Keynesian Phillips curves for country H and country F, respectively, where $\pi_{H,t} = \ln(P_{H,t}/P_{H,t-1})$ and $\pi^*_{F,t} = \ln(P^*_{F,t}/P^*_{F,t-1})$. They are derived by combining the log-linear approximation of the optimal price (A.31) with the log-linear approximation of (A.32) for each country separately. The parameter k is defined as

$$k = \frac{(1-\alpha\beta)(1-\alpha)}{\alpha}\frac{1}{1+\sigma\eta}. \tag{A.51}$$

Equation (A.48) is the log-linear approximation of the terms of trade (A.12), expressed in first differences. Equations (A.49) and (A.50) are the log-linear approximations of the country-specific CPIs (A.5) and (A.7), applying the law of one price and expressed in first differences.

A.3.2 Sticky-Price Equilibrium under Local-Currency Pricing

Under sticky prices and local-currency pricing, the system of equations is given by

$$E_t\hat{C}_{t+1} = \hat{C}_t + \frac{1}{\rho}\left(\hat{R}_t - E_t\pi_{t+1} + E_t\hat{\zeta}_{C,t+1} - \hat{\zeta}_{C,t}\right) \tag{A.52}$$

$$\hat{Q}_t = \rho\left(\hat{C}_t - \hat{C}^*_t\right) + \left(\hat{\zeta}^*_{C,t} - \hat{\zeta}_{C,t}\right) \tag{A.53}$$

$$E_t\Delta\hat{S}_{t+1} = \hat{R}_t - \hat{R}^*_t \tag{A.54}$$

$$\hat{Q}_t = (2a-1)\hat{T}_t + 2a\hat{\Delta}_t \tag{A.55}$$

$$\hat{Y}_{H,t} = 2a(1-a)\theta(\hat{T}_t + \hat{\Delta}_t) + a\hat{C}_t + (1-a)\hat{C}^*_t \tag{A.56}$$

$$\hat{Y}_{F,t} = -2a(1-a)\theta(\hat{T}_t + \hat{\Delta}_t) + (1-a)\hat{C}_t + a\hat{C}^*_t \tag{A.57}$$

$$\pi_{H,t} = (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) - (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}^{fb}_t + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] + k\hat{\mu}^H_t + \beta E_t \pi_{H,t+1} \tag{A.58}$$

$$\pi^*_{H,t} = (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}^{fb}_{H,t}\right) - (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}^{fb}_t + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] - k\hat{\Delta}_t + k\hat{\mu}^H_t + \beta E_t \pi^*_{H,t+1} \tag{A.59}$$

$$\pi_{F,t} = (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right) + (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}^{fb}_t + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] + k\hat{\Delta}_t + k\hat{\mu}^F_t + \beta E_t \pi_{F,t+1} \tag{A.60}$$

$$\pi^*_{F,t} = (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}^{fb}_{F,t}\right) + (1-a)k\left[2a(\rho\theta-1)\left(\hat{T}_t - \tilde{T}^{fb}_t + \hat{\Delta}_t\right) - \hat{\Delta}_t\right] + k\hat{\mu}^F_t + \beta E_t \pi^*_{F,t+1} \tag{A.61}$$

$$\hat{T}_t = \hat{T}_{t-1} + \pi_{F,t} - \pi^*_{H,t} - \Delta\hat{S}_t \tag{A.62}$$

$$\pi_t = a\pi_{H,t} + (1-a)\pi_{F,t} \tag{A.63}$$

$$\pi^*_t = (1-a)\pi^*_{H,t} + a\pi^*_{F,t} \tag{A.64}$$

$$\hat{\Delta}_t = \hat{\Delta}_{t-1} + \Delta S_t + \pi^*_{H,t} - \pi_{H,t}. \tag{A.65}$$

The Euler consumption equation, the risk-sharing condition, and the uncovered interest parity condition are identical to the case of producer-currency pricing. The real exchange rate is still linked to the terms of trade, but now also to the deviation from the law of one price (equation (A.55)). Since the countries are assumed to be symmetric, the deviation from the law of one price is identical across countries ($\hat{\Delta}_{H,t} = \hat{\Delta}_{F,t} = \hat{\Delta}_t$). The aggregate demand equations (A.56) and (A.57) as well as the four New Keynesian Phillips curves (A.58) through (A.61) contain the deviation from the law of one price as well. The terms-of-trade identity (A.62) and the definitions

of the CPI inflation rates (A.63) and (A.64) are different from the case of producer-currency pricing, since the law of one price does not hold under local-currency pricing. Finally, equation (A.65) is the log-linear approximation of the definition of the deviation from the law of one price (A.22), expressed in first differences.

A.3.3 Efficient Allocation

The first-best (fb) or efficient allocation describes the equilibrium in which prices are fully flexible, in which the law of one price holds, and in which markups are neutralized at all times with an appropriate subsidy ($\mu_t^i = 0$). This efficient allocation provides a useful benchmark in order to assess the welfare implications of the two currency regimes.

Accordingly, efficient output in each country is given by

$$\begin{aligned}(\rho+\eta)\tilde{Y}_{H,t}^{fb} =& 2a(1-a)(\rho\theta-1)\tilde{T}_t^{fb} \\ &- (1-a)\left(\hat{\zeta}_{C,t} - \hat{\zeta}_{C,t}^*\right) + \hat{\zeta}_{C,t} + \eta\hat{\zeta}_{Y,t} \\ (\rho+\eta)\tilde{Y}_{F,t}^{fb} =& -2a(1-a)(\rho\theta-1)\tilde{T}_t^{fb} \\ &+ (1-a)\left(\hat{\zeta}_{C,t} - \hat{\zeta}_{C,t}^*\right) + \hat{\zeta}_{C,t}^* + \eta\hat{\zeta}_{Y,t}^*.\end{aligned} \tag{A.66}$$

The efficient terms of trade can be written as

$$\begin{aligned}[4a(1-a)\rho\theta + (2a-1)^2]\tilde{T}_t^{fb} =& \rho\left(\tilde{Y}_{H,t}^{fb} - \tilde{Y}_{F,t}^{fb}\right) \\ &- (2a-1)\left(\hat{\zeta}_{C,t} - \hat{\zeta}_{C,t}^*\right).\end{aligned} \tag{A.67}$$

The first equation is obtained by combining the risk-sharing condition (A.41), equation (A.43), and the aggregate demand equation (A.44), all of which hold under flexible prices as well, with the log-linear approximation of the optimal-price equation (A.33). The second equation is derived completely analogously. The third equation is derived by subtracting the country-specific aggregate demand equations (A.44) and (A.45) from each other and by using the risk-sharing condition (A.41) and equation (A.43) to eliminate country-specific consumption and the real exchange rate.

Appendix B. Monetary Union Regime

The main difference of the monetary union regime compared with the flexible exchange rate regime, of course, is that the two countries share one currency and that the common monetary policy sets one union-wide nominal interest rate. Notwithstanding, the model structure is to a large extent identical (see, e.g., Benigno 2004). However, the type of local-currency pricing that was specified under the flexible exchange rate regime is impossible in a monetary union, since both countries share one currency. Accordingly, the law of one price is assumed to always hold under the monetary union regime.

Under flexible prices, monetary policy is neutral, so that real variables are only driven by fundamental shocks. Thus, the efficient allocation is independent of the currency regime. Therefore, the behavior of efficient output and the efficient terms of trade under the monetary union regime is also described by equations (A.66) and (A.67).

Under sticky prices, the system of equations is given by

$$E_t\hat{C}_{t+1} = \hat{C}_t + \frac{1}{\rho}\left(\hat{R}_t^{MU} - E_t\pi_{t+1} + E_t\hat{\zeta}_{C,t+1} - \hat{\zeta}_{C,t}\right) \tag{B.1}$$

$$\hat{Q}_t = \rho\left(\hat{C}_t - \hat{C}_t^*\right) + \left(\hat{\zeta}_{C,t}^* - \hat{\zeta}_{C,t}\right) \tag{B.2}$$

$$\hat{Q}_t = (2a-1)\hat{T}_t \tag{B.3}$$

$$\hat{Y}_{H,t} = 2a(1-a)\theta\hat{T}_t + a\hat{C}_t + (1-a)\hat{C}_t^* \tag{B.4}$$

$$\hat{Y}_{F,t} = -2a(1-a)\theta\hat{T}_t + (1-a)\hat{C}_t + a\hat{C}_t^* \tag{B.5}$$

$$\begin{aligned}\pi_{H,t} = {} & (\rho+\eta)k\left(\hat{Y}_{H,t} - \tilde{Y}_{H,t}^{fb}\right) - 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^H + \beta E_t\pi_{H,t+1}\end{aligned} \tag{B.6}$$

$$\begin{aligned}\pi_{F,t}^* = {} & (\rho+\eta)k\left(\hat{Y}_{F,t} - \tilde{Y}_{F,t}^{fb}\right) + 2a(1-a)(\rho\theta-1)k\left(\hat{T}_t - \tilde{T}_t^{fb}\right) \\ & + k\hat{\mu}_t^F + \beta E_t\pi_{F,t+1}^*\end{aligned} \tag{B.7}$$

$$\hat{T}_t = \hat{T}_{t-1} + \pi_{F,t}^* - \pi_{H,t} \tag{B.8}$$

$$\pi_t = a\pi_{H,t} + (1-a)\pi_{F,t}^* \tag{B.9}$$

$$\pi_t^* = (1-a)\pi_{H,t} + a\pi_{F,t}^*. \tag{B.10}$$

The Euler consumption equation (B.1) differs from the one under the flexible exchange rate regime only in that the nominal interest rate is given by the union-wide interest rate $\hat{R}_t^{MU}$. Nonetheless, real interest rates $\hat{R}_t^{MU} - E_t\pi_{t+1}$ are generally country specific, since CPI inflation rates usually differ across countries. Since the two countries form a monetary union, the uncovered interest parity condition is obsolete and the nominal exchange rate disappears from all relevant equations.

Appendix C. Robustness of Results

Figure C.1. Welfare Loss as a Function of the Inflation Coefficient (ϕ_π) for Different Combinations of Producer-Currency Pricing (PCP), Local-Currency Pricing (LCP), Productivity Shocks, and Cost-Push Shocks

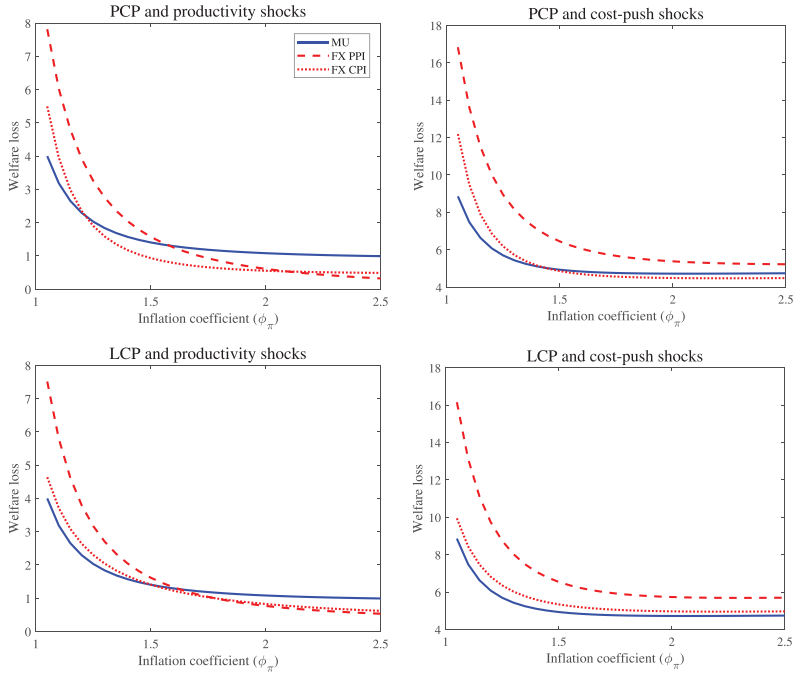
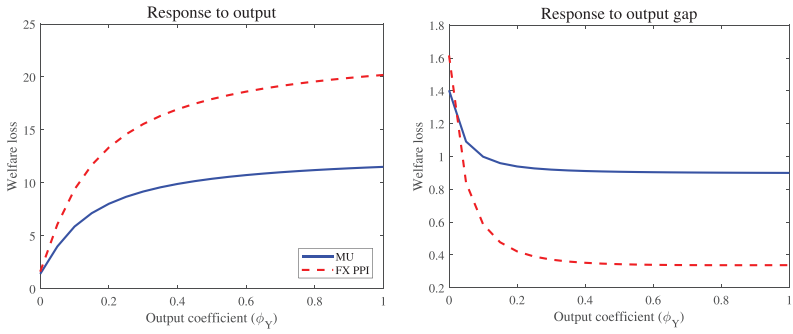
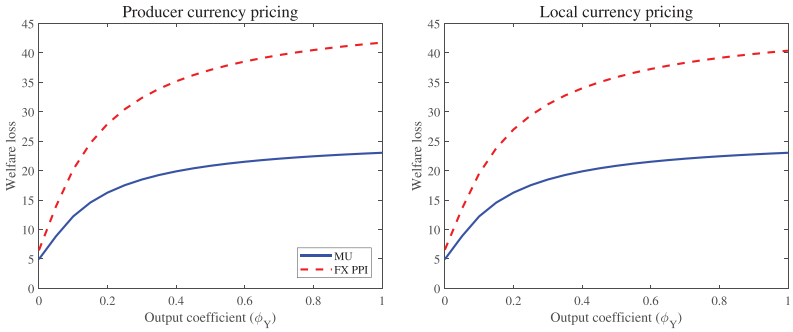


Figure C.2. Welfare Loss as a Function of the Output Coefficient (ϕ_Y) under Local-Currency Pricing and Productivity Shocks



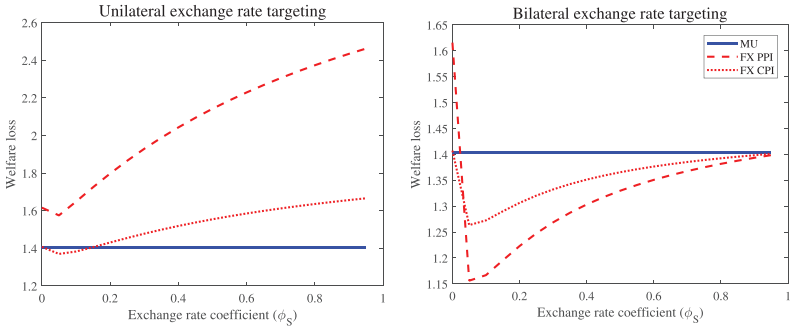
Notes: Left panel: Response to output ($\hat{Y}_t$). Right panel: Response to output gap ($\hat{Y}_t - \tilde{Y}_t^{fb}$).

Figure C.3. Welfare Loss as a Function of the Output Coefficient (ϕ_Y) under Cost-Push Shocks



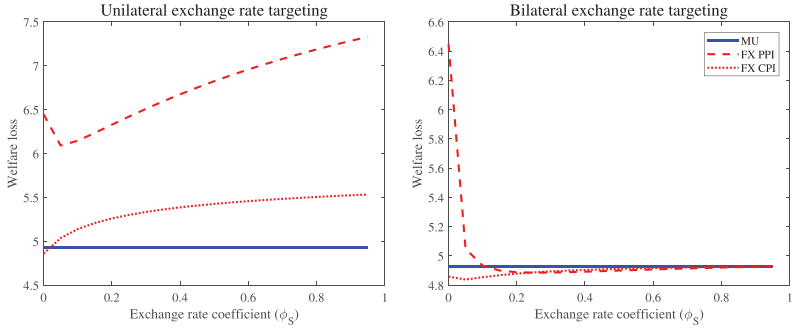
Notes: Left panel: Producer-currency pricing. Right panel: Local-currency pricing Under cost-push shocks, there is no difference between targeting output and targeting the output gap.

Figure C.4. Welfare Loss as a Function of the Exchange Rate Coefficient (ϕ_S) under Local-Currency Pricing and Productivity Shocks



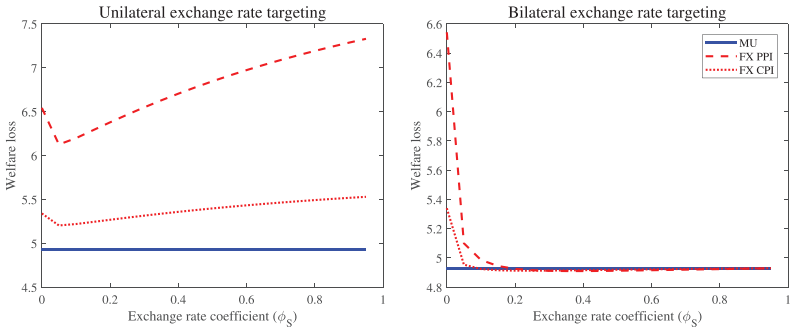
Notes: Left panel: Unilateral exchange rate targeting. Right panel: Bilateral exchange rate targeting.

Figure C.5. Welfare Loss as a Function of the Exchange Rate Coefficient (ϕ_S) under Producer-Currency Pricing and Cost-Push Shocks



Notes: Left panel: Unilateral exchange rate targeting. Right panel: Bilateral exchange rate targeting.

Figure C.6. Welfare Loss as a Function of the Exchange Rate Coefficient (ϕ_S) under Local-Currency Pricing and Cost-Push Shocks



Notes: Left panel: Unilateral exchange rate targeting. Right panel: Bilateral exchange rate targeting.

Figure C.7. Welfare Loss as a Function of the Interest Rate Smoothing Coefficient (ϕ_R) for Different Combinations of Producer-Currency Pricing (PCP), Local-Currency Pricing (LCP), Productivity Shocks, and Cost-Push Shocks

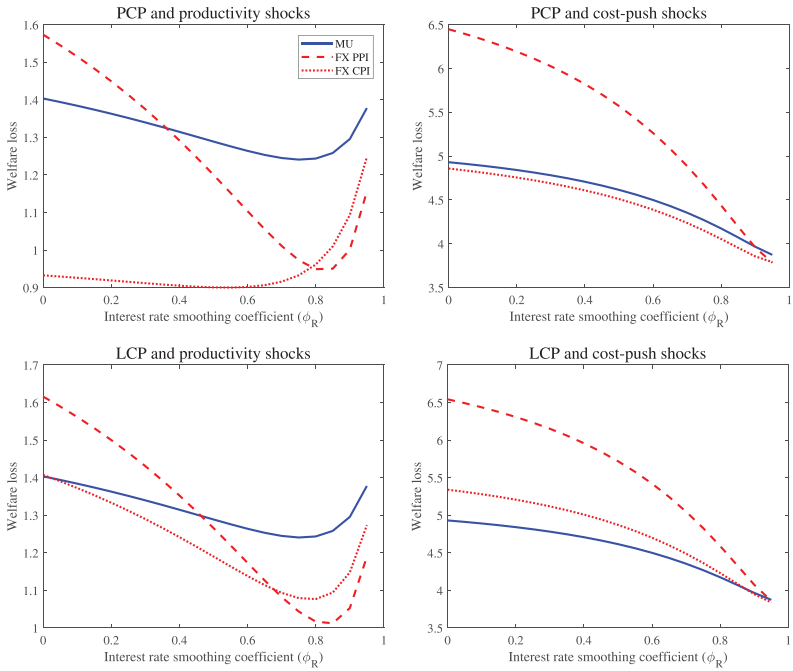
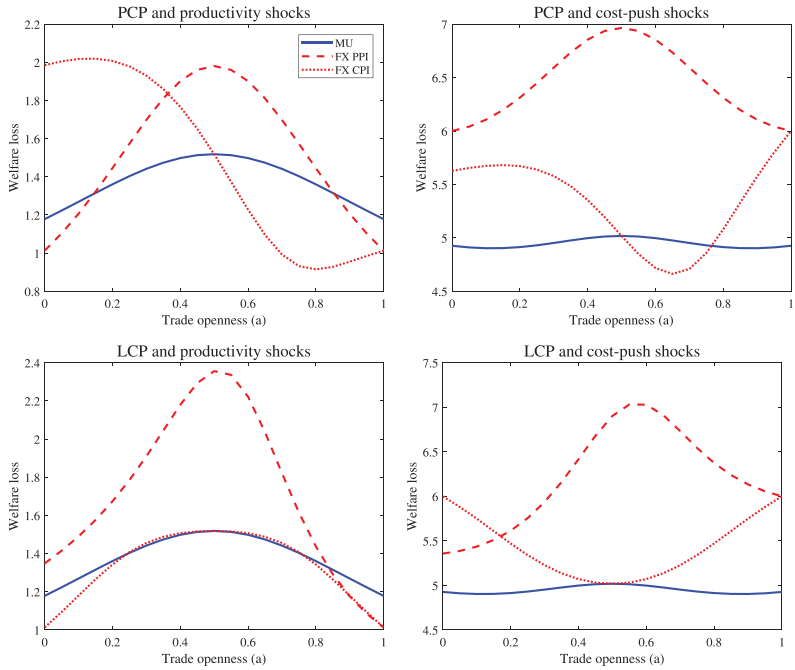


Figure C.8. Welfare Loss as a Function of the Trade Openness (a) for Different Combinations of Producer-Currency Pricing (PCP), Local-Currency Pricing (LCP), Productivity Shocks, and Cost-Push Shocks



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